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Unemployment in OECD countries

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Chapter 1

Introduction

The main aim of this thesis is to contribute to the understanding of the evolution of the unemployment rate in the OECD area. There is vast variation in unemployment rates both between countries and over time, and the explanations that have been given in the economic profession have also changed. The unemployment rates were very low in the beginning of the 1960s, but increased sharply in the 1970s and remained high throughout the next decade. The increase was first explained by shocks and fluctuations, i.e. mainly due to the two large oil price shocks in the 1970s. However, when unemployment remained high, this was explained by a slowdown in total factor productivity growth, i.e. a structural change, cf. Blanchard (2006a). In the late 1980s and early 1990s, labour market institutions were put forward as the dominating force behind the combination of high persistent unemployment and large differences between countries, see e.g. Layard and Nickell (1986), Layard et al. (1991) and OECD (1994). This view was supported by econometric evidence like in Nickell et al. (2005), where 55 percent of the increase in European unemployment over the period 1960s to 1990s was explained as due to changes in institutions. Around 2005, there was a gradual decline in the unemployment rates, and the unemployment problem received less attention. As stated in Boeri (2009), "Europe is no longer a continent of mass unemployment".

However, the financial crisis has led to a massive increase in unemployment rates in the OECD-area. This makes it appropriate to review our understanding of the unemployment problem. Is the "mainstream" view that variation in unemployment can, to a large extent, be explained by differences in labour market institutions, consistent with the evolution of unemployment up till the financial crises? And in the current downturn, can fiscal policy be used to combat unemployment?

This thesis can be seen as a contribution to the empirical literature of equilibrium unemployment. In this literature, explanations of changes in the unemployment rate are often decomposed into two parts; the factors that explain the long-run relationships of unemployment (the theory of equilibrium unemployment) and the factors that explain the fluctuations around equilibrium unemployment. I hope to contribute to both parts, and investigate both the effect of changes in institutions and government purchases.

This introduction summarises the issues that are discussed in this thesis, and comments on the main findings. I will also relate my work to current and previous research on the same issues. The starting point is a discussion of the theoretical framework for unemployment which is the foundation for the empirical analysis in the thesis. A brief discussion of the econometric method is given before the summary of each chapter.

There exist several theories of equilibrium unemployment and its fluctuations. One influential direction is laid out in the comprehensive book by Layard et al. (1991). In this approach, the equilibrium unemployment rate is determined by the intersection between the wage and the price curve in an unemployment-real wage diagram. The wage setting is represented by the outcome of a wage bargaining between the parties in the labour market. High unemployment weakens workers' position in the wage setting, implying that the wage curve is downward sloping. The price curve is the outcome of firms' price setting. Under constant returns to scale and mark-up pricing, the price curve is horizontal. However, the equilibrium unemployment rate is not interpretable as a constant given from nature. Equilibrium unemployment will increase with factors that shift either of the curves upwards. For instance, an increase in the wage pressure due to stronger union power or a higher tax rate will shift the wage curve upwards and raise equilibrium unemployment. Normally, the actual unemployment rate will not be equal to the equilibrium unemployment rate, temporary changes in global or government demand, or that the actual price level turned out to be lower or higher than expected (e.g. lower prices on import products due to the production from China) can push the economy out of equilibrium.

Layard et al. (2005) argue forcefully that the imperfections in the labor market are of such importance that every theory of unemployment has to put these imperfections in focus. This argument distinguishes their theory from the Neo-classical framework for the labour market, where workers maximize their utility depending on consumption and leisure, given the real wage. Employment is given by the intersection of individuals' optimal labour supply and firms' demand for labour at given wages. The theory assumes rational expectations and complete information. The theory implies that the unemployed workers have chosen not to work, since the outside option is better.

The wage-bargaining theory presented in Layard et al. (2005) is to a large extent consistent with the search-matching theory of unemployment. This theory was developed by the recent Nobel Price winners Peter Diamond, Dale Mortensen and Chris Pissarides. The theory explains unemployment by the fact that it usually takes time for employers and unemployed workers to match, and that the search process involves some costs. The same institutional variables that lead to high unemployment in the wage-bargaining theory also raise unemployment here, but the mechanism is somewhat different. For instance, stricter employment protection will make employers more reluctant to hire new workers, because it is more difficult to fire redundant workers in a recession. This increases the time spent on search, and thus also the equilibrium unemployment rate. A comprehensive review of these models is found in Pissarides (2000).

The search-matching model can also be extended to take into account changes in human capital; see Ljungqvist and Sargent (2008). In this paper, they explain why unemployment was lower in Europe than in the US before 1980, but lower in the US since then (disregarding the recent financial crisis). The key mechanism is the interaction between a turbulence parameter (that has increased over time), and employment protection and unemployment benefits that they claim to have been fairly constant over time. A combination of a low level of the turbulence parameter and strict employment protection leads to low frictional unemployment (short-term unemployment), because the flow of workers into unemployment is low when employment protection is strict and fluctuations are small. This was the situation in Europe before 1980, and hence resulted in low unemployment in this period. Since then, the world has become more risky, i.e. the economies have experienced larger fluctuations. Larger shocks imply that one person loses more human capital in case of an involuntary job loss, and there is an increase in structural unemployment (long-term unemployment). Ljungqvist and Sargent (2008) claim that both Europe and the United States are exposed to larger fluctuations after 1980, but that Europe has a higher structural level of unemployment due to stricter employment protection and higher unemployment benefits as compared to the United States.

The above theories do not always give any clear predictions about the effects of institutional variables on unemployment. One reason might, as already mentioned, be that the effect of the institution on unemployment is unstable over time, like the effect of strict employment protection in Ljungqvist and Sargent (2008). The effects of an institutional variable are also claimed to be dependent on the level of the institutional variable itself. For instance, a high degree of coordination among the wage setters may imply low wage claims if the level of coordination is above a certain level (Calmfors et al., 1988). This theory predicts high unemployment rates at a medium level of coordination, because at this level the wage setters are unable to influence the national level of unemployment, yet they have power to raise the wage above the free competition level. Low and high coordination levels result in lower unemployment levels: At a low coordination level, the wage setters have low bargaining power and hence, there is less wage pressure. With high coordination, the national unemployment level is affected by the outcome of the wage setting, and coordinated wage setters will therefore lower their wage claims to avoid high unemployment rates. A similar argument can be applied to union density. Wage setters that cover the whole economy will care about the unemployment levels in their wage setting, while unions that only cover parts of the economy will care less because they are unable to influence the aggregate unemployment level. Finally, unemployment benefits are normally assumed to increase unemployment because workers reduce their search effort, but if the existence of unemployment benefits causes workers to take more risk (benefits protect against the downside risk) by using time to find the right job, this can result in lower unemployment rates because better matching leads to higher productivity growth; see resent research e.g. Caliendo et al. (2009). The arguments made here illustrate that according to the theory, many institutional variables have ambiguous effects on unemployment. This shows that the theory needs to be subjected to empirical analysis, to find which is the dominating effect.

Panel data models for the equilibrium unemployment rate typically include variables representing the labour market institutions implied from the theory above; see Nickell et al. (2005), Bassanini and Duval (2006), Belot and van Ours (2004), Belot and van Ours (2001) and Blanchard and Wolfers (2000). The empirical papers have also included interaction terms between the labour market variables to account for the fact that the effect of institutions may depend on the existence of other institutions. For example, the unemployment rate might be lower if a country has a high degree of both coordination of wage setting and union coverage. The reason is that both these variables raise the unions' effect on the unemployment level. The unions will therefore lower their wage claims to avoid increasing unemployment rates. Other factors, like macroeconomic shocks and interaction between institutional variables and shocks, are also included in existing research. Previous literature documents a fairly robust positive correlation between unemployment benefits and unemployment, while the results in the literature of, for instance, the effect of labour market taxes and employment protection are less conclusive; see e.g. Belot and van Ours (2004).

The panel data approach has also been subject to extensive critique; see e.g. contributions by Baker et al. (2005), Blanchard (2006b), Berger and Everaert (2008) and Belot and van Ours (2004). The issues in this debate can roughly be divided into four parts: the size of the estimated coefficients, the variables in the empirical analysis, the chosen indexes, and the econometric method. However, it is difficult to evaluate how the various elements of the critique affect the overall validity of the results in empirical papers. The problem arises because key features like the time period, the econometric method, and the time series for the variables generally vary across papers. This is a recurrent problem in empirical literature, implying that it is difficult to compare new results with existing papers. When different data or econometric methods are used, it is hard to identify the source of the difference in results. This also makes it difficult to determine how robust the results are. In addition, the empirical specifications for unemployment do not in general distinguish or relate how the variables enter a dynamic specification from the static framework for equilibrium unemployment presented above. One strategy has been to use a simple first-order dynamics in the explanatory part of the model, and compensate by allowing for flexible dynamics in the residuals of the equations. On the other hand, basing the specification on heuristics alone also means that there is a gap between the underlying theory of equilibrium unemployment, which is static, and the dynamic specification used to estimate equilibrium rate unemployment.

Despite the critique presented above, panel data involves a clear advantage in analysing the reasons for differences in the unemployment rates due to the large variation between countries for the institutional variables that often exhibit little variation within countries. In table 1.1, the column with standard deviation shows that the between countries variation is larger for all variables except unemployment and unemployment benefits. The larger between variation provides a rationale for using the cross-country analysis, while the larger within variation for unemployment and unemployment benefit provides a rationale for using a time series analysis.

I have followed the panel data literature and included variables representing the labour market institutions in a final equation for unemployment. Compared to previous literature, I have extended the data set to the period 1960 to 2007 for 20 OECD countries. This gives more variation over the sample period, which covers both high and low levels of unemployment. Hopefully, this will result in more robust results, in particular compared with the studies over the period 1960 to 1995, which may have been dominated by the general increase in the unemployment rate over the same period.

It is appropriate to comment on the choice of estimation method, since this choice might influence the results. All estimated models in this thesis have some common structure; i.e. the models all include country-specific effects and have a dynamic structure. I have tried to follow the empirical guidelines for how to choose the right empirical specification for these types of models, especially for chapter 3 and 4, but the guidelines also have some consequences for how to interpret the main results in Nickell et al. (2005) (the estimation method is fixed effect (FE)) which is the topic of chapter 2.

Country-specific effects (one-way heterogeneity) are unobserved in our models, but if they are correlated with the other explanatory variables, the omission of fixed effects will lead to biased estimates of the other explanatory variables in the model (Baltagi, 2008). Even though it might be reasonable to assume such a correlation in our panel, this choice might still be questioned, as will be shown shortly.

A random effect (RE) model assumes there to be no correlation between the fixed effects and the explanatory variables included in the model. However, as previously mentioned, table 1.1 shows small within variation for some of the variables, like the coordination of wage setting. Using an RE model would leave more variation to estimate the coefficient of these variables. Since the RE approach assumes no correlation with the other regressors, this additional orthogonality condition can be tested by a Hausman test,

Table 1.1: Labour market institutions and unemployment. Time period 1960 to 2007

Variable		Mean	Std. Dev.	Min	Max	Observations
Unemployment	overall	5.37	3.56	0.00	19.11	N = 1019
	between		2.00	1.58	9.40	n = 20
	within		2.98	-2.51	16.93	T-bar = 50.95
Empl. Protection	overall	2.13	1.22	0.00	4.19	N = 1020
	between		1.19	0.11	3.98	n = 20
	within		0.38	0.67	2.94	T = 51
Unempl. benefits	overall	0.42	0.21	0.00	0.89	N = 1020
	between		0.14	0.21	0.66	n = 20
	within		0.15	-0.16	0.87	T = 51
Benefit duration	overall	0.45	0.35	0.00	1.04	N = 1020
	between		0.31	0.04	1.02	n = 20
	within		0.17	-0.09	0.87	T = 51
Union density	overall	0.41	0.19	0.08	0.84	N = 890
	between		0.18	0.14	0.76	n = 20
	within		0.08	0.08	0.70	T = 44.5
Coordination	overall	3.40	1.24	1.00	5.00	N = 960
	between		1.09	1.00	4.90	n = 20
	within		0.64	0.86	5.46	T = 48
Tax	overall	0.44	0.13	0.16	0.79	N = 1012
	between		0.12	0.21	0.64	n = 20
	within		0.07	0.14	0.62	T = 50.6

see Wooldridge (2002, Ch. 10.7.). The results of such a test in chapter 4 show that the Hausman test rejects the RE model at a 6 percent level. Thus, we have chosen a fixed effect (FE) model where we allow for correlation between the country-specific effects and the other regressors.

Ordinary least squares on a model with both fixed effects and a lagged endogenous variable will, in general, result in biased and inconsistent estimates, see Baltagi (2008). However, the long time series mitigate the fixed effect bias, cf. Judson and Owen (1999). The alternative to the FE approach is to transform the model to first differences, which eliminates the fixed effect, and use instruments for the endogenous first difference of the lagged endogenous variables, see (Baltagi, 2008). This method leads to unbiased and consistent estimates, and one such approach is the Arelleano-Bond method. However, the long time series augments the number of available instruments and these must be reduced, see Roodman (2009). In spite of some suggestions in this literature of how to reduce the number of available instruments, there does not exists a clear path. On this data set, cf. an extensive discussion of the instrumental variable estimation in chapter 3, it turned out to be difficult to achieve robust results with this estimation method. This aspect, and the fact that the fixed effect bias is small with long time series, have led us to the conclusion that the FE-estimation method is preferable. The small bias on long time series is also a reason why this estimation method is not used in the analysis of the replication of Nickell et al. (2005) in chapter 2.

The existence of an equilibrium level for unemployment requires that there is no unit root in the unemployment rate. The unit root of unemployment with a data generating process up to lag three, subtracted country-specific means, is rejected by two formal tests (Levin et al. (2002) and Im et al. (2003)) in chapter 3.

Now let us turn to the thesis itself, which includes three chapters in addition to the introduction. The starting point is a replication of the results in Nickell et al. (2005), including an investigation of how robust the effect of labour market institutions is to data revisions and time series extensions. A simulation of the main model for unemployment in Nickell et al. (2005) reveals that the predictive power of the model is weak, and that the model underpredicts the unemployment rate for 17 out of 20 countries in the post sample period. Despite this, the analysis also shows that countries that changed their institutions in an 'employment-friendly' way experienced lower unemployment rates in the post sample period, compared to countries that changed their institutions in the opposite direction.

Chapter 3 re-visits the question of the role of institutions. Compared to chapter 2, the econometric model is a dynamic version of the wage bargaining theory as specified in Layard et al. (2005). The dynamic specification implies some notable differences from the existing empirical literature: First, the third-order unemployment dynamics in the final equation of unemployment is a consequence of the structural model, which depends on both the number of equations and the order of dynamics in those equations. This is an extension of earlier papers that use a first-order dynamics (or second-order dynamics if the residuals in addition are of first order). A second result is that the underlying theory has implications for the signs and the magnitude of the coefficients of the lags of unemployment, which can be confirmed or refuted by estimation. Third, the labour market variables should enter with lag one and two according to the derivation. Fourth, the formal derivation of the dynamic unemployment equation also makes it clear that there is no logical or *a priori* reason why the equilibrium unemployment rate cannot be a function of other factors than the labour market institutions. Our empirical results support the chosen dynamic specification. Furthermore, they show that temporary changes in the economic environment have had a larger impact on the evolution of the unemployment rate than institutions.

The final chapter of the thesis expands the demand side of the model in chapter 3, and includes both government purchases and changes in export demand. The latter variable is probably more sensitive to aggregate macroeconomic trends, and can be viewed as a control for the business cycle or as a control for macroeconomic shocks that hit all economies in the panel at the same time. We find that an increase in government purchases leads to a clear reduction in unemployment, even if the effect varies across OECD countries.

1.1 Summary of chapter 2

Using panel data for twenty countries from 1960 to 1995, Nickell et al. (2005) find that labour market institutions explain most of the variation in OECD unemployment. They find that 55 percent of the increase in European unemployment are due to changes in institutions, where changes in the unemployment benefit system and taxes are the main contributions. The specified model includes year dummies, country-specific dummies and time trends, to avoid that the included variables are distorted by omitted variables with trends for each country or global shocks. They have also specified some variables that are supposed to capture temporary changes in the economic environment, like changes in labour demand shocks and changes in total factor productivity.

There are several reasons for re-assessing the paper by Nickell et al. (2005). First, the results are strong, explaining the bulk of the variation in unemployment by variation in institutions. It is also noteworthy to obtain homogenous effects of institutions for all countries in the panel. Second, the strand of research focussing on the link between labour market institutions and unemployment has been very influential. The results of this research have been interpreted as supporting the recommendation from international organizations like OECD to countries for how they should change their institutions in order to reduce unemployment rates, e.g. reduce the level of unemployment benefits.

The large effect of changes in institutions on unemployment found by Nickell et al. (2005) is a natural starting point for investigating the role of institutions. The first question addressed in the chapter is whether the results in Nickell et al. (2005) could be used to forecast the evolution of unemployment in the OECD countries. Ex post, we now perfectly know the evolution of unemployment and the explanatory variables; thus, we can test whether their explanations are consistent with the subsequent evolution of the unemployment rate. A dynamic simulation of the main model in Nickell et al. (2005) from 1995 to 2007 shows strong underprediction for 17 of the 20 countries, while unemployment is only overpredicted for one country.

I then explore three possible explanations for the underprediction; evolution of shocks, change in the data generating process and possible misspecifications of the model in Nickell et al. (2005). In this investigation, I use the methodology derived in Nickell et al. (2005), to ensure that the variation in results are not due to a new estimation technique or method.

The first obvious candidate for an explanation of the underprediction is that the shocks that are included as explanatory variables in the empirical model have evolved differently in the post sample period. The results of a simulation with variation in the shock variable and a simulation where the shock variables are set to zero in the post sample period, show nearly no effect of the shocks in the extended period, with Japan and Italy being the only exceptions.

The second candidate is to investigate if the data generating process has changed, i.e. if the link between institutions and unemployment has changed, in the post sample period. The model in Nickell et al. (2005) is reestimated on the revised and extended data set, and the results show that the size of the coefficients changes quite substantially in both time periods. I also repeat their analysis of calculating the long-run effect of changes in institutions for the European countries. I find that changes in institutions now account for 76 percent of the total change in unemployment from the 1960s to 2002-2007, up from 41 percent in my replication of their results over the shorter time period 1990 to 1995. At face value, this might suggest that institutions have become more important. However, the interpretation is less clear cut. First, the larger share reflects that unemployment increased less over the longer period, so there is less to explain. In addition, a dynamic simulation does not give a better fit with the reestimated coefficients than with original ones, even for the extended sample period, suggesting that changes in the coefficient values is not the key explanation for the underprediction.

I then investigate the model dynamics to detect if this is the cause of the underprediction in the post sample period on the original data set. A dynamic simulation of the full model, when the error term is explicitly taken into account, reveals that the model is non-stationary for some countries. This is verified by the roots of the 2nd order differential equation that is implied by the model in Nickell et al. (2005). The results suggest a reconsideration of the dynamic specification of the model. Especially since the underlying solution to the specified model implies a 2nd order dynamics in the unemployment rate. However, this is not the main cause for the underprediction of the unemployment rate, as most of the countries have a stable solution to the model. Instead, it turns out that the underprediction of the unemployment rates is largely driven by the dynamic specification of the model, where the combination of the large coefficient for the lagged unemployment rate, the trend and the fixed effects, implies a tendency for unemployment to diverge in one direction or the other. This implies that forecasting a stationary time series as the unemployment rate is impossible.

In light of the severe underprediction of unemployment, it seems worthwhile to explore the link between institutions and unemployment in isolation, ignoring the other parts of the model. Using the estimated coefficients, the change in labour market institutions would predict that average unemployment increases by 1.3 percentage points over the period 1995 to the average over the period 2002 to 2007, while actual unemployment fell by 2.3 percentage points in the same period. Once more, this suggests that the model does not capture the effects well. However, if one takes country dispersion into account, a different picture emerges. There is a clear tendency that countries which have changed their institutions in an "employment-friendly" way, like Denmark and Finland, have experienced a larger reduction in unemployment than countries that have changed their institutions in the opposite direction like Germany and Portugal. This indicates that labour market institutions affect unemployment in the direction found by Nickell et al. (2005).

1.1.1 Future research

The main methodological contribution from this analysis is that one should be very careful in modelling time trends in unemployment models. The analysis shows that the underprediction is largely driven by the model dynamics and that the accounting exercise in

Nickell et al. (2005) is not suited to analysing how much of the variation in unemployment that can be attributed to changes in institutions in the post-sample period. In addition, it should be kept in mind that specifying dynamics in the disturbance term might affect the overall stability of the system. Equilibrium theory requires stable solutions for unemployment.

Despite of this methodological issue, the substantial contribution of Nickell et al. (2005), i.e. that labour market institutions affect unemployment in the direction found by Nickell et al. (2005), is to a large degree still valid. There is a clear tendency that countries which have changed their institutions in an “employment-friendly” way, like Denmark and Finland, have experienced a larger reduction in unemployment than the countries that have changed their institutions in the opposite direction like Germany and Portugal. On the other hand, small effects of the included shocks illustrate that it is not these shocks that should be the focus of future research.

1.2 Summary of chapter 3

The theoretical literature reviewed above is static, while both static and dynamic specifications of the unemployment equations are used in the empirical literature reviewed above. The dynamics is reasonable given that there are adjustment lags in the manifold of economic, administrative and political decisions that jointly determine the rate of unemployment.

The existing studies rely on heuristics to motivate the dynamic specification of the econometric panel data model. Heuristics gives the empirical researcher considerable freedom to choose a specification that fits the data well. One strategy has been to use simple first-order dynamics in the regression and compensate by allowing flexible dynamics in the residuals of the equations. On the other hand, basing the specification on heuristics alone also means that there is a gap between the underlying theory of equilibrium unemployment, which is static, and the dynamic specification used to estimate equilibrium rate unemployment.

In the third chapter of the thesis, we estimate the quantitative importance of labour market institutions for equilibrium unemployment in the OECD. Compared to existing literature, the econometric model is based on the solution of a dynamic macroeconomic model, which includes structural equations for wage and price setting. We also use a sample with more variation in unemployment and institutions, and a higher order dynamics in the final equation for unemployment. Finally, we incorporate objectively and automatically selected indicators for structural breaks in the unemployment rate. The argument for this latter assumption is that the effect of changes in institutional variables on unemployment is likely to be gradual, and are modelled by relatively long lags in accordance with theory. Therefore, we interpret the intermittent but large changes in the unemployment rate from one year to another to be due to other factors than institutions, like extraneous or domestic demand shocks, changes in households' preferences for work and leisure or changes in pro-or counter-cyclical economic policies. We call these changes structural breaks. We have used two statistical methods for detecting such shocks, one of which leads to fewer breaks that the other.

We find that institutional variables have statistical significance, but that these variables account for relatively little of the overall change in the OECD average unemployment rate. For instance, the most robust effect of labour market institutions is changes in the benefit replacement ratio. Based on the long-run estimates from our main equation, we find that if the benefit replacement ratio is lowered by 20 percent, from the OECD average in 2007, the unemployment rate will decrease by 0.8 percentage points. On the other hand, we find that the absence of large negative shocks to the economy has been more important for the reduction in the actual average unemployment rate from the 1990s and up until the resent financial crises. The inclusion of structural breaks that capture location shifts in the distributions for the unemployment rates turns out to be important for our estimate of the equilibrium rate. If we do not correct for these structural breaks, the equilibrium rate is simulated to almost 6.2 percent, while the lowest adjusted estimate is 4.3 percent. However, comparing the simulation of the two models with structural breaks shows that the model with more breaks illustrates a larger gap between time varying and constant institutions. This could illustrate the importance of controlling for other factors influencing unemployment in order to achieve the true effect of institutions.

In terms of modelling methodology, this paper illustrates the importance of a dynamic specification of the panel data model for the rate of unemployment. We show that a reduced lag structure on the autoregressive coefficients increases the residual autocorrelation which might be a sign of misspecification. The chosen dynamic specification is theoretically derived and has the status of a final equation of a system consisting of equations for wage and price setting and an equation of unemployment as a function of the real exchange rate. The theoretical derivation gives a priori assumptions regarding the magnitude of the autoregressive coefficients. The magnitude is confirmed by the empirical evidence. On the other hand, our results also show that the exact lag structure of the institutional variables is of minor importance for capturing the effects of labour market institutions on unemployment.

1.2.1 Future research

In spite of the advantages of using panel data, there are also some difficulties. Even with substantial effort, empirical work will never be able to model the world perfectly, but researchers aim at controlling for large outliers and other factors that are not specified in the model and that might bias the estimates of the other variables included in the model. For instance, if an econometrician is aware of special historical events that have had an substantial impact on the economy and the endogenous variable of interest, like wars or the breakdown of the former Soviet Union, the empirical model will try to correct for these events by including dummies. A dummy will exclude this event from having an impact on the endogenous variable and all explanatory variables in the model. In this way, researchers try to avoid that their estimates of the included variables in the model are biased from these events.

Panel data studies might be especially vulnerable to such problems because of heterogenous events. When modelling several countries, detecting all such events might be even more difficult than when modelling only one country. In general, panel data analysis that models macro variables such as the unemployment rates includes time- and country-specific dummies to capture country-specific effects and macroeconomic shocks that are not explicitly modelled.

One problem with this approach is that we lose many degrees of freedom in estimating both country-specific time trends and common or country-specific time dummies in long time series. We have used a different approach to account for events in a more objective way than, for instance, reading the history books and trying to subjectively detect which special events that should be accounted for in estimating the final equation of unemployment. The method is a statistical method for detecting special events in the time series of interest, known as "impulse saturation" and "large outlier" approach. The properties of this class of automatic model selection procedures using *Autometrics* are discussed in Castle et al. (2010) and Hendry and Mizon (2010). The method is by now well known in time series analysis, but not yet in panel data analysis. The paper in chapter 3 must therefore be seen as a first attempt to use these statistical methods developed for time series analysis on a panel.

We have followed the time series analysis to account for the breaks country by country. However, as our knowledge has improved, we have discovered that it is possible to first apply panel data techniques to transform our variable of interest, the unemployment rate, to avoid country-specific variation, and stack the data set with all countries in the panel. Then, the above methods can be directly used on this stacked data set.

The results of this process are somewhat different from our first method, but still the results are not too different from the country by country method used in this thesis. In future work, this line could be of considerable interest for panel data econometricians, since they could, in an objective way, detect whether they should control for special events by starting out by assuming that every year is a special year, and then reducing the number of year dummies in this statistical and objective way. If the time series are long, this will also increase the degrees of freedom if the alternative is one time dummy per year.

The long time series could gain by testing the stability of the parameter values as in normal time series approach. This could be a preferable way of investigating how the parameters change when adding a country, and also when adding one observation to the regression. For example, this could possibly reveal changes in the effect of employment protection as captured by Ljungqvist and Sargent (2008).

1.3 Summary of chapter 4

In the final chapter of my thesis, we expand the demand side of the model with government purchases. This is an interesting exercise also in light of the recent financial crisis. Most OECD countries used the fiscal policy extensively to combat the crisis by stimulating the economy. More recently, fiscal policy has been reversed in many countries. The large changes in policy raise several key questions in relation to the effect on unemployment; in particular whether fiscal policy measures can be used to combat increasing unemployment, and if fiscal tightening is likely to lead to persistent high unemployment.

Whether and possibly to what extent fiscal policy should be used to stabilize the economy is a question subject to a great deal of political controversy. In contrast, more concrete questions, like how will an increase in government spending affect unemployment, should be less controversial. However, there is no consensus in the literature on the effect of fiscal policy. There is now a rapid growth in the literature, and hopefully a more consensus view may emerge.

We test the quantitative importance of government purchases for the evolution of unemployment in the OECD. The analysis is built on the preferred empirical specification in chapter 3, and adds the change in government purchases and an export market indicator as explanatory variables. Compared to earlier studies, we use a sample with more variation in unemployment and institutions.

We find that increased government purchases lead to lower unemployment; the point estimate is that an increase equal to one percent of GDP reduces unemployment by 0.2 percentage points in the same year, and increases to 0.25 percentage points after one year, to then gradually vanish over the following decade. The effect is greater in downturns than in booms, and also greater under a fixed exchange rate regime than under a floating regime.

One methodological problem in the analysis is that government purchases might be endogenous, in the sense that fiscal policy decisions clearly depend on the state of the economy. This might be the case even if the endogeneity problem is likely to be less severe for government purchases as compared to, for instance, transfers, since government purchases are not directly linked to the state of the economy. We address the endogeneity problem by using instruments and including omitted variables. This is a different approach as compared to previous studies which also try to address this problem; see Perotti (2007), Beetsma and Giuliodori (2010) and Hall (2009) for recent reviews. In our view, also the alternative methods have their weaknesses, and our method should be considered as complementary to the other studies mentioned. We show that our main results are robust to our methods of addressing the endogeneity problem, even if the instrumental variable approach indicates that the effect of government purchases is downward biased in the fixed effect estimation.

1.3.1 Future research

In light of the recent financial crises and the large changes in fiscal policy, it would be particulary interesting to extend the data set used in chapter 4 to cover also this period, and explore if the effect of government purchases remains robust. However, it would be difficult to disentangle the effect of fiscal policy from other shocks.

Another natural extension of the chapter would be to explore the effect of taxes as a part of fiscal policy, and not a labour market institution as is done in the current analysis.

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Chapter 2

The robustness of empirical models for unemployment. A review of Nickell et al. (2005).

Victoria Sparrman

Abstract In an influential study, Nickell, Nunziata and Ochel (2005) find that institutions explain most of the variation in OECD unemployment, using panel data for 20 countries from 1960 to 1995. The importance of Nickell et al. (2005)'s conclusions has spurred a lively debate, and several authors have criticized their findings. This paper reassesses the main findings in Nickell et al. (2005), benefitting from the inclusion of twelve additional years of data. A dynamic simulation of their main unemployment equation shows that unemployment is severely underpredicted in the post sample period for 17 of the 20 countries, while it is only overpredicted for one country. The analysis shows that the underprediction is largely driven by the model dynamics, and that the accounting exercise in Nickell et al. (2005) is not suited to analyze how much of the variation in unemployment that can be attributed to changes in institutions in the post-sample period. However, there is a clear tendency that countries which have changed their institutions in an "employment-friendly" way, like Denmark and Finland have experienced a larger reduction in unemployment than the countries that have changed their institutions in the opposite direction like Germany and Portugal. This indicates that labour market institutions affect unemployment in the direction found by Nickell et al. (2005).

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2.1 Introduction

A number of recent papers have tried to explain the evolution of unemployment in the OECD area, based on an equilibrium unemployment framework. One of the most influential is Nickell, Nunziata and Ochel (2005) NNO hereafter. They find that the development in labor market institutions can account for 55 percent of the increase in European unemployment for the period 1960 to 1995. In their analysis, the effect of institutions on unemployment is the same for all the 20 OECD countries in the panel.

There are several reasons to re-asses the paper by NNO. First, the results are strong, explaining the bulk of the variation in unemployment by variation in institutions. It is also noteworthy to obtain homogenous effects of institutions for all the countries in the panel. Second, the strand of research focussing on the link between labour market institutions and unemployment has been very influential. NNO is a major, recent contribution to this literature, spurring a lively debate and receiving 461 references in Google Scholar. This approach has strongly influenced the recommendations from international organizations such as the OECD on how countries should organize their economies. The results of this strand of research have been interpreted as a recommendation to countries for how they should change their institutions in order to avoid or reduce unemployment. According to the findings of this literature, reducing institutional variables like the benefit replacement ratio, employment protection, union density and taxes would lower the unemployment rate considerably.

This paper evaluates the results in NNO over the extended sample 1995 to 2007. I ask the question of whether the NNO results could be used to forecast the evolution of unemployment in the OECD countries, if one were able to predict perfectly the evolution of the explanatory variables. To this end, I undertake a dynamic simulation of the model for unemployment with the original estimated coefficients. Dynamic simulation from 1995 to 2007 shows strong underprediction for 17 of the 20 countries, while unemployment is only overpredicted for one country. The difference between simulation and actual unemployment rate for the period 1995 to 2007 motivates a closer look at the empirical model in NNO. I also take a closer look at NNO's finding that 55 percent of the increase in unemployment is caused by changes in institutions in Europe in the sample period. I explore whether this result survives data revisions within the sample period, and what the method gives for the post-sample period.

This paper is not the first to discuss the results and method in NNO, see for instance contributions by Baker et al. (2005), Blanchard (2006), Berger and Everaert (2008) and in Belot and van Ours (2004). The issues in this debate can roughly be divided in four: the size of the estimated coefficients, the variables in the empirical analysis, the chosen indexes, and the method. However, in spite of the extensive literature, it is difficult to evaluate how the various elements of the critique affect the overall validity of the

NNO results. The problem arises because key features like the time period, the method, and the time series for the variables generally vary across papers. This is a recurrent problem in empirical literature, implying that it is difficult to compare new results with existing papers. When different data or methods are used, it is hard to identify the source of the difference in results. This also makes it difficult to determine which parts of the original results that are still valid, and which parts that are not.

In this paper every source of variation in results is separated by looking at one source at the time: revision of time series and extension of the time period. In addition, a replication of the original results in NNO ensures that the same method is used throughout the paper.

The results from the extended time period show that the data generating process has changed somewhat, but the change in estimated coefficients has not led to a markedly better fit of the model in the extended time period. Institutions explain a larger share of the increase in unemployment in Europe than over the shorter period until 1995, but this is because the actual increase in unemployment up to 2007 is lower. The results for how much of the increase in unemployment that can be explained by institutions are not convincing and the underprediction of the unemployment rate in the post sample period is still unexplained.

The empirical specification of the model in NNO is evaluated further in section 2.5. A dynamic simulation where the specified error term is taken into account in the simulation illustrates that the estimated model has a non stationary solution for three of the countries; Japan, Netherland and New Zealand. However, also for the other countries a simulation of only the dynamic part of the equation, i.e. the lagged unemployment rate, the trend and the fixed effects, yields severe instability. Stable dynamics is essential to be able to predict a stationary time series such as unemployment. Most likely, this is the source of the underprediction.

One possible objection to the present analysis is that the empirical model of NNO was developed to explain the evolution of unemployment in the period 1960 to 1995, and to explore the link between institutions and unemployment, but not to predict unemployment. However, the high impact of the policy recommendations clearly show that the results in this and similar studies have been regarded as being of general validity. Moreover, if the empirical model of NNO captures the crude features of the data generating process, and this data generating process has been fairly stable over time, then one would expect the model also to be able to perform well in a post-sample dynamic simulation, given the correct values of the explanatory variables. A failure in the post-sample simulation would suggest that either the model explains unemployment behavior less well than the in-sample results indicate, or that the unemployment behavior has changed over time. Both conclusions would be of considerable interest, as well as motivate further research.

In light of the severe underprediction of unemployment, it seems worthwhile to explore the link between institutions and unemployment in isolation, ignoring the other parts of the model. Using the estimated coefficients, the change in labour market institutions would predict that average unemployment increase by 1.3 percentage points over the period 1995 to the average over period 2002 to 2007, while actual unemployment fell by 2.3 percentage points in the same period. Again, this suggest that the model does not capture the effects well. However, if one takes country dispersion into account, a different picture emerges.

The paper is organized as follows. First I present the development in actual unemployment in the post-sample period, some related literature and the empirical model in NNO , which is used throughout the paper in section 2.2. Then, in section 2.3 I show that this model underpredicts the unemployment rate in 17 of the 20 countries in the panel. The model in NNO is reestimated in section 2.4, and the results show that the data generating process has changed somewhat, but the reestimation does not lead to a markedly better fit. In section 2.5 I explore the dynamic specification of the model and find that it is unstable for most countries. The link between institutions and unemployment in the post-sample period are evaluated from a different angle in section 2.6. Section 2.7 concludes. Appendix 2.A describes the construction of the data. Appendix 2.B presents some additional results that do not change the main picture of the prediction results in section 2.3. The replication of the model in NNO and some additional results to section 2.4 and 2.5 are presented in appendix 2.C.

2.2 Background

Before turning to the model specification, it is interesting to look at the actual development in the unemployment rates in the sample period available, i.e. 1960 to 2007. Then, the theoretical framework for explaining the evolution in unemployment as given in NNO is presented.

2.2.1 The evolution of unemployment in the OECD area

The unemployment rate in the OECD countries changed substantially in the period 1960 to 2007, see table 1. For instance, Switzerland has a relatively low but increasing unemployment throughout the period, while in Ireland and Spain on the other hand the unemployment rate is much more volatile. Germany and Japan experience a steady increase in the unemployment rate over time, ending on a fairly high unemployment. The unweighed average unemployment rate (see bottom row) was very low at the beginning of the period, but increased sharply and peaked at the decades 1980 and 1990. Then unemployment fell slightly in the last period 2002 to 2007.

Figure 1 also shows the unemployment rate for all countries in the sample in the period 1960 to 2007. We observe that the rise in unemployment in the early 1970s went together

Table 1: The unweighed average of unemployment. Revised and extended data set. Percent

Country	1960-64	1965-72	1973-79	1980-87	1988-95	1996-01	2002-07
Australia	1.75	1.79	4.66	7.70	8.41	7.33	5.31
Austria	1.70	1.42	1.38	3.25	4.89	5.49	5.67
Belgium	1.48	1.48	4.23	9.61	8.05	8.34	8.05
Canada	6.00	4.76	6.98	9.84	9.53	8.11	6.92
Denmark	1.07	1.04	3.56	6.48	7.50	5.00	4.62
Finland	1.41	2.41	4.14	5.17	10.85	11.54	8.34
France	1.18	1.95	3.71	7.67	9.10	9.66	8.48
Germany	0.69	0.86	3.06	6.56	6.94	8.31	9.29
Ireland	5.32	5.82	8.08	14.05	14.68	7.30	4.47
Italy	3.46	4.17	4.87	7.96	9.91	10.81	7.71
Japan	1.34	1.24	1.84	2.52	2.46	4.22	4.62
Netherlands	0.57	1.26	3.57	8.28	6.60	4.29	4.03
New Zealand	0.08	0.29	0.74	3.95	8.14	6.37	4.12
Norway	1.71	1.53	1.74	2.44	5.13	3.69	3.89
Portugal	2.46	3.91	5.63	8.23	5.48	5.23	6.90
Spain	1.78	2.31	4.04	14.51	15.00	13.61	9.76
Sweden	2.11	2.61	2.62	3.59	6.22	9.06	6.92
Switzerland	0.03	0.01	0.29	0.63	2.24	3.30	3.99
UK	2.79	3.40	4.81	10.44	8.77	6.31	5.10
United States	5.72	4.47	6.51	7.75	6.16	4.63	5.27
Total	2.14	2.34	3.82	7.03	7.80	7.13	6.17

with an increase in the dispersion across the countries in the sample. After 1995, both the average unemployment rate and the variation in the unemployment rates across countries have decreased. In appendix 2.A, figure A1 displays the evolution of unemployment for various groups of countries.

2.2.2 The empirical specification as given in NNO

The approach of NNO is based on an equilibrium unemployment framework. In the short run actual unemployment may deviate from equilibrium unemployment due to shocks, but unemployment eventually returns to its equilibrium level. The equilibrium unemployment theory could be based on several different micro founded theories of unemployment, like wage bargaining, efficiency wages, or search and matching theories, see e.g Layard et al. (1991) and Pissarides (2000).

The empirical model for unemployment in table 5 in NNO is specified by a simultaneous system that consists of the following two equations:

$$\begin{aligned} U_{it} \;=\;& \theta U_{i,t-1} + \beta_1 EPL_{it} + \beta_2 BRR_{it} + \beta_3 (BD_{it} - \overline{BD}) * (BRR_{it} - \overline{BRR}) \\ & + \beta_4 (UDNET_{it} - UDNET_{i,t-1}) + \beta_5 CO_{it} + \beta_6 (CO_{it} - \overline{CO}) * (UDNET_{it} - \overline{UDNET}) \\ & + \beta_7 TW_{it} + \beta_8 (CO_{it} - \overline{CO}) * (TW_{it} - \overline{TW}) + \alpha_1 LDS_{it} + \alpha_2 DPROD\_hp_{it} \\ & + \alpha_3 TTS_{it} + \alpha_4 D2.MS_{it} + \alpha_5 RIRL_{it} + \gamma 1 t + \gamma 2_i + \gamma 3_i t + v_{it} \qquad (1) \end{aligned}$$

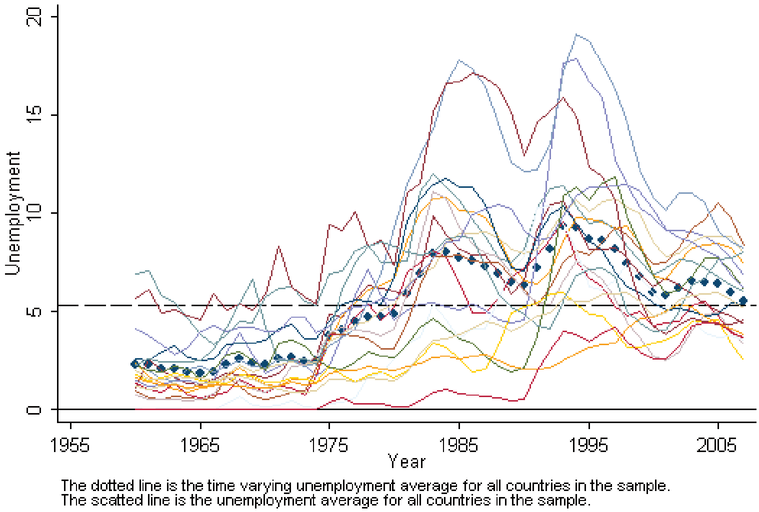

Figure 1: Actual unemployment. Revised and extended data set over the years 1960 to 2007. Percent

and

$$v_{it} \quad = \quad \rho_i v_{i,t-1} + \epsilon_{it} \tag{2}$$

The institutional variables and the interaction terms among these variables in the first part of equation (1) determines the equilibrium unemployment level. The institutional variables are indexes for employment protection (EPL), benefit replacement ratio (BRR), benefit duration (BD), union density ($UDNET$), tax rate (TW) and coordination of wage setting (CO). The interactions are: benefit duration and benefit replacement ratio, coordination in wage setting and union density, and coordination and tax rate. The interaction terms are measured as deviation from the variable mean. The second part of equation (1) consists of five variables meant to capture deviations from the equilibrium unemployment rate over the business cycle, the shocks. The shock variables are; labour demand (lds), total factor productivity ($DPROD_{hp}$) , import prices (TTS), money supply ($D2.MS$) and real interest rate ($RIRL$). ($U_{i,t-1}$) is the unemployment in the previous period. Finally, the heterogenous part of the model is captured by unobserved country and time specific shifts in the intercepts ($\gamma 1$ and $\gamma 2$), a country specific trend ($\gamma 3_i * t$) and a country specific autoregressive error term (v_{it}). The latter is defined in equation (2), where ρ_i is the country specific auto regressive coefficient and ϵ_{it} is white noise.

The coefficients in equation (1) are estimated with feasible general least squares, where the ρ_i in the error term is estimated simultaneously for every country i by a iteration process using panel data for 20 OECD countries over the period 1960-1995.

The results in appendix 2.C, in model A in table C1, and are in line with what we would expect from the equilibrium theory of unemployment: Higher taxes and more generous unemployment benefits increase unemployment, while more coordination in wage setting leads to lower unemployment.

NNO chose to evaluate the empirical model and the effect of institutions by a dynamic simulation of the unemployment rate, disregarding the error term by setting the expected value of the error term in equation (2) to zero in every period and for every country. They claim that the close visual similarity between the simulation and the actual unemployment rate over the estimation period illustrates that the model explains the data well.

They claim, based on the dynamic simulation keeping institutions fixed at their 1960s level (i.e. the unweighted mean over the period 1960 to 1969), that institutions account for 55 percent of the increase in European unemployment in the period 1960s to the early 1990s, measured as the increase in the unweighted mean of unemployment from the 1960s to the period 1990-1995. Changes in the benefit system are the most important, contributing 39 percent. Increases in labour taxes generate 26 percent, shifts in the union variables are responsible for 19 percent, and movements in employment protection law contribute 16 percent.

2.3 Forecast of the unemployment rate by using the NNO model in the post-sample period 1995 to 2007

This section re-assesses the findings of NNO, benefitting from twelve additional years of data. Specifically, I explore to what extent the empirical model of NNO is able to forecast the subsequent post-sample evolution of unemployment, given that we now in general know the correct values of the explanatory variables.

I evaluate the model by use of static and dynamic simulation of unemployment. The model does very well in a static simulation, see appendix 2.B. However, as the static simulation is conditional on the lagged unemployment rate, which plays a large role in the model, the static simulation may not give the right impression about the model's out of sample explanatory power. I will thus focus on the dynamic simulation, which is also used by NNO. In a dynamic simulation the simulated value of unemployment in period $T + t$ is used to forecast unemployment in period $T + t + 1$, see Clements and Hendry (1998, Ch. 2.7).

In general, the sample period is extended by using the time series for the institutional variables available up to 2003, except for taxes that are available up to 2007, and the time series for the macro variables that are available up to 2007. Note that even if most of the institutional variables only are available up to 2001, 2002 and 2003, this is not a major problem for predicting unemployment up to 2007 by this model. This is because changes in institutions are slow and changes take a long time to materialize in the unemployment rate. With a estimated coefficient of the lagged dependent variable of 0.86, 47 percent of the full effect of changes in institutions remains after 5 years ($0.86^5 = 0.47$). This means that most of the changes in institutions up to 2003 have materialized in unemployment by 2007, while changes in institutions after 2003 will have little impact before 2007. See appendix 2.A for details regarding the extended and revised data set. The time dummies for new years are set to zero in the simulations. The time trend is prolonged and the estimated coefficients from the original data set are used in the simulations.

Formally, the dynamic simulation of unemployment in the current period can be written as $\tilde{U}^d_{i,T+t} = E(U_{i,T+t}|\tilde{U}^d_{i,T+t-1}, \mathbf{X}_{i,T+t}, \hat{\beta})$, where $\tilde{U}^d_{i,T+t-1}$ is the dynamic simulated unemployment rate for country i in the previous period, $\mathbf{X}_{i,T+t}$ is a vector that contains all explanatory variables for country i in period $T+t$ [1] and $\hat{\beta}$ is a vector with all the within sample period estimated coefficients as given in NNO[2]. The error term is set to zero in every period in this simulation, but the simulations of unemployment follow the same pattern if the error term is equal to the last estimated value of the country specific error term (the results are not reported here).

The dynamic simulation results are shown in figure 2. The figure reveals a large disparity between the dynamic simulation and the actual unemployment rate for most countries in the post sample period. The simulated unemployment rate is lower than the actual unemployment rate for 17 countries, i.e. Australia, Austria, Belgium, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Sweden, Switzerland, United Kingdom and United States. The simulated unemployment rate is substantially higher than the actual rate for Spain. The model simulates well for only two countries, Finland and Ireland.

If the empirical model specified by NNO explains the development in unemployment, the difference between the simulated and the actual value of the unemployment can be caused by the development in the institutional variables, the specified shock variables, the error term, or how the trend and time dummies are prolonged.

The discrepancy between the simulated and actual unemployment rate remains even if the trend or the time dummies are prolonged in various ways. In the following, some examples will be given. The estimated trend is negative for most countries in the sample, exceptions are Finland, Ireland, New Zealand and Spain, but the simulated unemployment rate follows a similar pattern if the trend variable is prolonged by the last value of the

[1]The vector contains mostly actual values of the explanatory variables, but it also contains some predictions for some of the institutions in the period 2001-07, see appendix 2.A for details regarding time periods for the different variables.

[2]The country specific dummies are taken from the replication in section 2.4.1

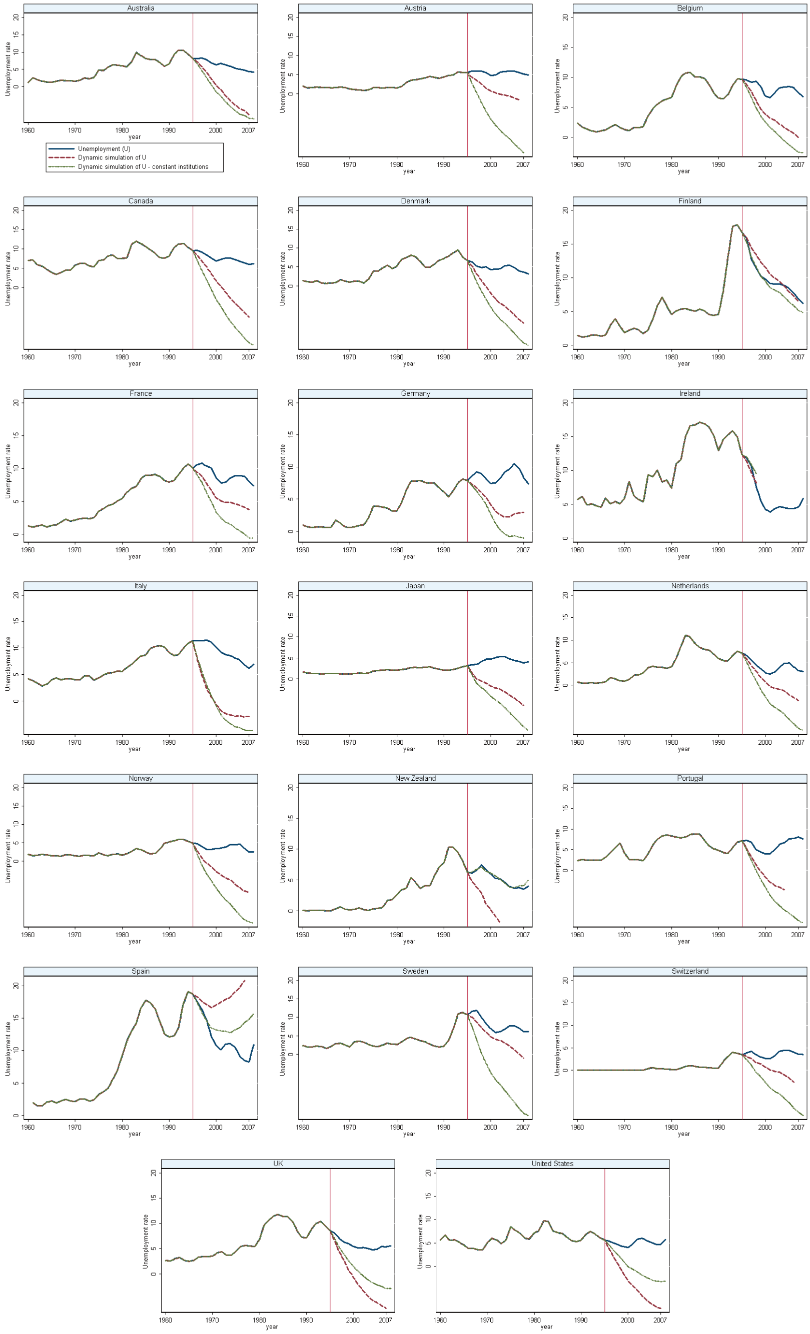

Figure 2: Actual unemployment and dynamic simulation of the unemployment rate with and without time-varying institutions. Estimated coefficients on the original data set are used in simulation, see appendix 2.C, model A in table C1. Institutions are constant. Percent

trend in 1995 or by the average value of the trend. The latter gives the largest change in simulations and is shown in appendix 2.C, figure B3. The simulations show that in this specification the predictive power of the empirical model improves somewhat for most countries in the sample, i.e. the simulated value of unemployment by the model is closer to actual unemployment in the post sample period, but the improvement is small compared with the distance between actual and simulated unemployment. The prediction of the model improved also in Sweden, but the model now overpredicts the change in unemployment. The predictive power worsens for New Zealand, Germany and United Kingdom. The simulations of the unemployment rate follow a similar pattern even if the time series are prolonged by the estimated average of the time dummies (results omitted for space considerations).

The specified shock variables have a small impact on the simulations, compare to the simulations in figure 2 and B4, where the latter is found in appendix 2.B. Thus, the underprediction is not caused by the shocks included in the NNO model.

On the other hand figure 2 illustrates a large effect of changes in institutions on the development of the unemployment rates. More specifically, the figure displays the actual unemployment rate, the dynamic simulation described above and a dynamic simulation where the values of the institutional variable are kept fixed at their 1995 level. For 14 of the 20 countries, incorporating the changes in the institutions leads to a reduction in the distance between actual and simulated unemployment rate. This is the case for Australia, Austria, Belgium, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, Norway, Portugal, Sweden, and Switzerland. For Ireland, there is no difference between the simulated unemployment rate with constant and time-varying institutions. However, for Finland, New Zealand, Spain, United Kingdom and United States, including the variation in institutions increases the difference between actual unemployment and the simulated unemployment rate. There is large effect of changes in institutions on unemployment even if we modify the trend, the time dummies or the shocks as described above.

The dynamic simulation of unemployment presented in figure 2 motivates a further investigation of the model in NNO post sample. The predictive power of the NNO model is weak, but changes in institutions still contributes to explain some of the development in unemployment for 11 of the 15 European countries. One is intrigued to know what the reason for the underprediction is. Has the data generating process, i.e. the link between the explanatory variables and unemployment changed after 1995? Or were there weaknesses in the empirical specification already in the original time period? The underprediction may also raise questions about the findings that changes in institutions explain the bulk of the increase in unemployment in Europe. Thus, it becomes important also to extend this analysis to the longer sample period. This is the topic of the following sections, where every source of variation in results is separated by looking at one source at the time: revision of time series and extension of the time period. Finally, the model dynamics and institutions effect on unemployment are investigated from a different angle.

2.4 Empirical investigation

In this section the data generating process is investigated. I also explore if these changes affect the explanatory power of the changes in institutional variables on unemployment.

The starting point in section 2.4.1 is an attempt to replicate the main results in NNO on the original data. Then, the empirical model in NNO is estimated on the revised data on two sample periods 1960 to 1995 and 1960 to 2007 in section 2.4.2.

2.4.1 Replication of model 1 in table 5 in NNO

By using the original data set to replicate the empirical model presented in table 5 in NNO [3] ensures that any differences in the results in NNO and in the results I will present in the following, are due to changes in data revisions, definitions or sample length. The estimation procedures and results in NNO are described in section 2.2.

I find that the estimated coefficients exactly replicates the coefficient values in NNO, cf. appendix 2.C, model A in table C1. In addition, a detailed visual inspection of the dynamic simulation on the original data set with and without time-varying institutions are also the same as in NNO. The replicated simulations are presented in appendix 2.C, figure C1.

As the lagged unemployment rate enters among the regressors, the long run effect of a change in institutions differs from the short run effect. To find the exact long run effect of changes in institutions in the period 1960s to 1990s, NNO compare the outcome of a dynamic simulation with the actual development of the institutional variables with dynamic simulation where the institutional variables are kept fixed at their 1960s values as described in section 2.2. To replicate their analysis, I also compare the outcome of a dynamic simulation with changing institutions with the outcome of a dynamic simulation where institutions are fixed at the 1960s values. The results which differ somewhat from those of NNO, are presented in table 2 and are commented below. However, due to the discrepancy between NNO and my results, I also calculate the long run effect by a second method. I compare the long run multiplier also used by Nickell and Nunziata (2002) p.19. This method intends to capture the permanent effect of a change in institutions by the standard formulae; the value of the estimated coefficient of the variable that represents the institution multiplied by the change in the same variable and divided by one minus the value of the estimated coefficient of the lagged unemployment rate. For instance, the long run multiplier for the change in employment protection is equal to $0.47 * \Delta epl/(1-\theta)$. Note that the two methods should not give the same results. The long run multiplier

[3] The data is received from Luca Nunziata.

is intended to capture the total effect of a change in the variables that represent the institutions over an infinite time period, and not the exact effect of these variables to unemployment in the limited time period 1960 to 1995. If institutional variables that have a negative impact on unemployment have changed late in the sample period, the full effects will not yet have materialized within the sample period, and hence the formula will exaggerate the effect of institutions on unemployment.

Table 2: Effects of institutions on unemployment in Europe. Original data set and the estimated coefficients in appendix 2.C, model A in table C1

Actual change in the unemployment rate and overall effects of institutions:							
	Actual Unemployment (U)	Dynamic simulation of U, all institutions change:	Dynamic simulation of U, all institutions constant:	Contribution of institutions to U: Dynamic simulation; percentage points	share of total[d]	Long run multiplier[c]; percentage points	share of total[d]
$\bar{U}^a_{60-69}$	2.06	2.09	2.01				
$\bar{U}^b_{90-95}$	8.86	9.16	4.60				
$\bar{U}^b_{90-95} - \bar{U}^a_{60-69}$	6.80	7.07	2.59	4.48	0.66	5.23	0.77

Decomposition to specific institutions:						
	Dynamic simulation of U, all institutions change: $\bar{U}^b_{90-95} - \bar{U}^a_{60-69}$	Dynamic simulation of U, one type of institution constant: $\bar{U}^b_{90-95} - \bar{U}^a_{60-69}$	Contribution of specific institutions to U: Dynamic simulation; percentage points	share of total[e]	Long run multiplier[c]; percentage points	share of total[e]
Constant:						
Benefits	7.07	3.77	3.30	0.74	3.73	0.71
Unions	7.07	7.16	-0.09	-0.02	0.47	0.09
Taxes	7.07	6.15	0.91	0.20	0.66	0.13
Empl. protection	7.07	6.71	0.36	0.08	0.37	0.07
Sum			4.48	1.00	5.23	1.00

a) $\overline{U}_{60-69}$ is the unweighted average of the unemployment rate, simulated or actual, in the period 1960 to 1969.
b) $\overline{U}_{90-95}$ is the unweighted average of the unemployment rate, simulated or actual, in the period 1990 to 1995.
c) The long run multiplier is calculated by the use of formula; $\frac{\hat{\beta}*\Delta X}{1-\theta}$ with the actual change in the specific institution (X).
d) Share of total increase in unemployment in Europe.
e) Share of total increase in unemployment explained by institutions in Europe.

The upper panel in table 2 shows that the actual unweighted average unemployment rate increased with 6.8 percentage points, from 2.06 to 8.86, in the period 1960s to 1990s ($\overline{U}_{90-95} - \overline{U}_{60-69}$), this is the same as in NNO. Table 2 upper panel also shows that in the dynamic simulation with time-varying institutions, unemployment increases by 7.07 percentage points from 1960s to 1990s ($\overline{U}_{90-95} - \overline{U}_{60-69}$), while with constant institutions, unemployment increases by 2.59 percentage points. Thus, changes in institutions can explain 4.48 percentage points of the actual 6.8 percentage points increase in the unemployment rate over the sample period, or 66 percent. The long run multiplier gives a somewhat higher share, as changes in institutions account for 77 percent of the increase in unemployment in Europe. The dynamic simulation method of calculating the contribution from institutions to unemployment is higher than the reported results in NNO, but lower than the long run multiplier method.

The lower panel in table 2 decomposes the contribution by the different types of institutions. In the dynamic simulation where benefits are kept constant, while the other institutional variable vary over the sample, unemployment increased by 3.30 percentage points. When all institutions change, unemployment increased by 7.07 percentage points, thus the increase in the benefit replacement ratio contributes to $7.07 - 3.77 = 3.30$ of the total increase by dynamic simulation of 4.48. Of this, the increase in replacement ratio accounts for 74 percent, union variables (union density, coordination and interaction between the two union variables) -2 percent, tax variables (tax rate and the interaction between the tax rate and coordination) 20 percent and employment protection 8 percent. The decomposition effects are different from the contributions reported in NNO. Especially, the effect of benefits is higher than the one reported in NNO. The long run multiplier shows that the increase in the benefit system accounts for 71 percent, unions 9 percent (coordination, union density and the interaction between these two variables), taxes 13 percent (tax rate and interaction between coordination and taxes) and employment protection legislation 7 percent of the actual increase in the unemployment rate. The decomposition on types of institutions is also here quite different from that of NNO, especially the effect of the changes in the benefit replacement ratio which accounts for 71 percentage points here versus 39 percentage points in NNO.

In sum: The replication of regression results has been successful with respect to obtaining the exact same coefficients values and a close visual similarity in the dynamic simulation with constant and time varying institutions as compared with NNO. The long term effects are somewhat different, in spite of my attempt to use the same method as they use. In the following I use the approach described above to compare the long run effects of institutions using the extended data set.

2.4.2 Estimation on the revised and extended data set

This section explores the effect of data revisions and an extended sample length on the results derived in the previous section. To isolate the effect of data revision, I first replicate the results on the original sample period, and then extend the sample length.

The economic variables are revised after the publication of NNO. Generally the data revisions are small. A comparison of the unemployment rate in table 1 with a similar table in NNO show no clear pattern of the unemployment rates being revised up or down since NNO presented their work. The largest increase in the unemployment rate was in Sweden, where it is revised up with more than 1 percentage point for several years in the sample. This revision was from a very low level implying a large relative percentage increase. The unemployment rates are revised down in Belgium, Denmark, Netherland, Spain and Switzerland throughout the period. The data revision ranges from $-0,17$ to $-4,59$ percentage points, with the largest decrease in Spain. Some of the decrease in the unemployment rate for Spain could be explained by that the time series used here is from OECD (2008a) while NNO used a time series from International Labour Organization (ILO). However, the revision in percentage is larger for instance for Belgium and Switzerland in some years. See appendix 2.A for data details.

The model in section 2.4.1 and data revisions

Figure 3 and table 3 present the estimated coefficients of the model in NNO on the revised data set over the time period from 1960 to 1995 [4]. The left panel of figure 3 illustrates the estimated coefficients on the revised data set divided by the replicated coefficients in appendix 2.C, model A in table C1. The value is equal to one if the coefficient is equal in both estimations. In general, the figure illustrates a reduced direct effect of institutions to unemployment on the revised data set, as most of the normalized coefficient values are smaller than one. In addition, a comparison of the left panel in table 3 with table C1, model A, where the latter is found in appendix 2.C, not only shows that the effects of benefit replacement ratio and coordination to unemployment are reduced, but they are also outside the original confidence interval. However, these estimated coefficients are still positive and significant. The estimated direct effects of employment protection, benefit duration and union density have changed sign, but are not significant. All the coefficients of the interaction terms are reduced. Two of the shocks, money supply and labour demand, have a stronger effect on unemployment than previously reported. The real interest rate, productivity and import price shock have a smaller effect than previously reported. The effect of the real interest rate is no longer significant.

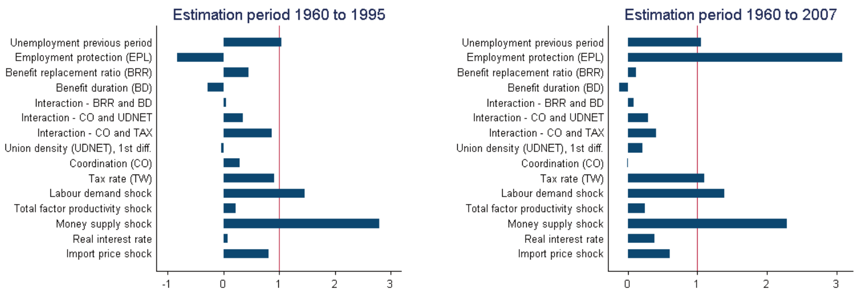

Figure 3: Estimated coefficients of equation (1) on the revised and extended data in table 3 divided by the original estimated coefficients in appendix 2.C, in model A in table C1 (value 1 means identical estimate)

The changes in the values of the coefficients indicate that the empirical specification is sensitive to small changes in the data set, i.e. to the revisions of the time series. This may reduce the relevance of their results to other samples. Note also that the persistence of unemployment in the previous period has increased in the empirical model from 0.86 to 0.90.

[4]Note however, that the revised data set includes 33 more observations than the original data set, because missing observations due to missing data for shock variables are included by setting the shocks to zero. The extra observations does not affect the main picture. In addition, to obtain convergence in the specified empirical model on the extended time period the time dummies are pooled for four more years, i.e. in the period 1960 to 1969. The change does not lead to large changes in the estimated coefficients on the original data set, compare model A and B in table C1, where the latter is found in appendix 2.C.

Table 3: Estimation of model 1 in table 5 in NNO on the extended and revised data set

	1960 to 2007				1960 to 1995			
	Coef.	t-value	min95	max95	Coef.	t-value	min95	max95
Unemployment previous period	0.90	65.64	0.88	0.93	0.89	45.66	0.86	0.93
Employment protection (EPL)	0.46	2.84	0.14	0.78	-0.13	-0.54	-0.58	0.33
Benefit replacement ratio (BRR)	0.27	1.08	-0.22	0.77	0.98	2.91	0.32	1.64
Benefit duration (BD)	-0.06	-0.45	-0.29	0.18	-0.13	-0.77	-0.47	0.21
Interaction - BRR and BD	0.33	0.54	-0.88	1.55	0.19	0.23	-1.42	1.79
Interaction - CO and UDNET	-2.08	-3.76	-3.16	-0.99	-2.42	-3.19	-3.91	-0.93
Interaction - CO and TAX	-1.41	-1.56	-3.17	0.36	-2.98	-2.48	-5.34	-0.62
Union density, 1st diff. ($\Delta UDNET$)	1.52	0.93	-1.67	4.70	-0.32	-0.18	-3.82	3.19
Coordination (CO)	0.00	0.02	-0.37	0.37	-0.29	-1.12	-0.81	0.22
Tax rate (TW)	1.67	2.52	0.37	2.97	1.37	1.70	-0.21	2.95
Labour demand shock	-32.74	-16.67	-36.59	-28.89	-34.54	-14.84	-39.10	-29.98
Total factor productivity shock	-4.44	-5.16	-6.13	-2.76	-3.87	-4.06	-5.73	-2.00
Money supply shock	0.53	2.27	0.07	0.98	0.64	2.48	0.13	1.15
Real interest rate	0.69	0.90	-0.81	2.19	0.13	0.14	-1.64	1.90
Import price shock	3.50	2.41	0.65	6.34	4.70	2.82	1.43	7.97
Total numb. of obs.	853				633			
Time periods Numb. groups	45	20			33	20		
Log likelihood	-822				-656			
χ^2 of all exogenous variables	666757				565092			

The aggregate long run effects of institutions on unemployment are presented in the upper panel of table 4. The results from the dynamic simulation with and without time-varying institutions show that changes in institutions can explain 2.78 percentage points of the 6.79 percentage points actual increase in the unemployment rate in Europe in the time period 1960s to 1990s, i.e. 41 percent. The results from the long run multiplier effect of changes in institutions to unemployment is 55 percent.

Table 4: Effect of institutions on unemployment in Europe. Coefficients in table 3. The revised and extended data set

Actual change in the unemployment rate and overall effects of institutions:

	Actual Unemployment (U)	Dynamic simulation of U, all institutions change	Dynamic simulation of U, all institutions constant	Contribution of institutions to U: Dynamic simulation, percentage points	Contribution of institutions to U: Dynamic simulation, share of total[e]	Contribution of institutions to U: Long run multiplier[d], percentage points	Contribution of institutions to U: Long run multiplier[d], share of total[e]
$\bar{U}^a_{60-69}$	2.04	2.04	2.03				
$\bar{U}^b_{90-95}$	8.84	8.40	5.59				
$\Delta\bar{U}_{60s-90s} = \bar{U}^b_{90-95} - \bar{U}^a_{60-69}$	6.79	6.35	3.57	2.78	0.41	3.74	0.55
$\bar{U}^c_{00-07}$	6.36	6.54	3.24				
$\Delta\bar{U}_{60s-00s} = \bar{U}^c_{00-07} - \bar{U}^a_{60-69}$	4.32	4.49	1.22	3.28	0.76	3.27	0.76

Decomposition to specific institutions::

	Dynamic simulation of U, all institutions change: $\Delta\bar{U}_{60s-90s}$	Dynamic simulation of U, all institutions change: $\Delta\bar{U}_{60s-00s}$	Dynamic simulation of U, one type of institution constant: $\Delta\bar{U}_{60s-90s}$	Dynamic simulation of U, one type of institution constant: $\Delta\bar{U}_{60s-00s}$	Contribution of specific institutions to U: Dynamic simulation; 1960s-1990s percentage points	Dynamic simulation; 1960s-1990s share of total[f]	Dynamic simulation; 1960s-2000s percentage points	Dynamic simulation; 1960s-2000s share of total[f]	Long run multiplier; 1960s-1990s percentage points	Long run multiplier; 1960s-1990s share of total[f]	Long run multiplier; 1960s-2000s percentage points	Long run multiplier; 1960s-2000s share of total[f]
Constant:												
Benefits	5.99	4.11	6.35	4.49	0.36	0.13	0.39	0.12	2.24	0.60	0.39	0.12
Unions	6.32	4.12	6.35	4.49	0.03	0.01	0.37	0.12	-0.79	-0.21	0.40	0.12
Taxes	5.08	2.80	6.35	4.49	1.27	0.46	1.69	0.53	2.36	0.63	2.05	0.63
Empl. protection	5.28	3.74	6.35	4.49	1.07	0.39	0.76	0.24	-0.07	-0.02	0.43	0.13
Sum					2.74	1.00	3.21	1.00	3.74	1.00	3.27	1.00

a) $\overline{U}_{60-69}$ is the unweighted average of the unemployment rate, simulated or actual, in the period 1960 to 1969.

b) $\overline{U}_{90-95}$ is the unweighted average of the unemployment rate, simulated or actual, in the period 1990 to 1995.

c) $\overline{U}_{00-07}$ is the unweighted average of the unemployment rate, simulated or actual, in the period 2000 to 2007.

d) The long run multiplier is calculated by the use of formula; $\frac{\hat{\beta} * \Delta X}{1-\theta}$ with the actual change in the specific institution (X).

e) Share of total increase in unemployment in Europe.

f) Share of total increase in unemployment explained by institutions in Europe.

Thus, with revised data, the change in institutions explains a smaller share of the increase in unemployment than with the original data set. Surprisingly, as the data now differs, the results are more in line with those reported by NNO.

The lower panel of table 4 shows the contribution of institutions on unemployment. The dynamic simulation shows that the largest contributors to the increase in the simulated unemployment rate are the increase in taxes which accounts for 46 percent and the increase in employment protection which accounts for 39 percent on the revised data. The main contributors to increased unemployment by the long run multiplier method are benefits and taxes, which contribute with 60 percent and 63 percent. Changes in union variables and employment protection decrease the unemployment rate in the period 1960s to 1990s with −21 percent and −2 percent.

The model in section 2.4.1 and extended sample length

We now turn to the effect of extending the sample period. The coefficient estimates on the extended time period are also divided by the original coefficients in the right panel of figure 3. We observe that all the institutional variables except for benefit duration have the same sign as the original coefficients. The direct effect of employment protection and taxes have increased. The effect of benefit replacement ratio, the union variables and all the interaction terms are reduced.

Table 3 summarises the estimated coefficients and the related t-statistics. As seen from the table, employment protection, taxes and the interaction term between the union variables have a significant positive effect on unemployment in the extended time period. For employment protection the significance may reflect that increased variation due to the extended time period have made it easier to capture the effect on unemployment. The benefit and union variables no longer have a significant effect on the unemployment rate. By comparing the results in table 3 and the original estimates in appendix 2.C, in model A table C1, the coefficient estimates of employment protection, benefit replacement ratio and taxes are the only among the institutional variables that are within the original 95 percent confidence interval. The coefficients of labour demand shock and money supply have increased and the latter is now also significant. The coefficients for the other shocks have decreased, but the effects are still significant.

It is interesting to look at the static and dynamic simulations of the model to investigate whether the values of the estimated coefficients from the revised data set result in a better fit regarding changes in the actual unemployment rate and the contribution of changes in institutions on unemployment.

In light of the considerable changes in the coefficient estimates, one might expect that the model would result in a bad fit of unemployment, at least for the period 1960 to 1995. However, it seems that the new coefficients in a static simulation simulate unemployment well over that period, see appendix 2.C, figure C2. One possible explanation is that the model is quite flexible because of time dummies and country specific trends, in the sense that the model can capture actual unemployment also with the new values of the institutional and shock variables estimated on the revised data set. However, as mentioned in section 2.3, the static simulation is conditional on the lagged unemployment rate, which plays a large role in the simulation implying that static simulation is not a powerful way to test the model. Thus, I also undertake a dynamic simulation.

The dynamic simulation of the estimated coefficients from table 3 is presented in figure 4. The overall picture is that the estimated model on the revised and extended data set simulates unemployment well for most countries in the sample, but not so good when the unemployment rate increases sharply at one point in time, cf. for instance Belgium, Canada, Finland, Sweden and Switzerland. However, this seems to be the case also with the dynamic simulation based on the original estimated coefficients. The simulations of the two empirical models are compared in figure 5. As seen from the figures, the original empirical model in NNO has a better visual fit for 10 countries as compared with the empirical model on the extended and revised data set. The new model has a better visual fit for 6 countries.

The effect of institutions is illustrated in figure 4. We observe that the simulation using time varying institutions tracks actual unemployment quite well for most countries, while the simulation with constant institutions quite often is far off the evolution of actual unemployment. This suggests that changes in institutions still are important in explaining the development of unemployment.

The long run effect of institution is summarised in table 4. Based on the dynamic simulation with constant and time-varying institutions, the upper panel shows that the actual unemployment rate increased by 4.32 percentage points, from 2.04 to 6.36, in the period 1960s to the average of the 2000 to 2007 ($\overline{U}_{00-07} - \overline{U}_{60-69}$). Changes in institutions account for 76 percent of the increase in unemployment in Europe from 1960s to the 2000s ($\overline{U}_{00-07} - \overline{U}_{60-69}$). By using the formula for the long run multiplier, institutions account for the same percentage share as the dynamic simulation, 76 percent. According to this exercise, institutions explain a larger share of the increase in unemployment than for the period 1960s to 1990s, and also a larger share than in NNO. Note however, that the reason is that actual unemployment increased less. The increase in unemployment that according to this analysis is due to changing institutions is actually lower for the longer period until the 2000s.

The decomposed effect of institutions on unemployment is presented in the lower panel of table 4. The change in benefits account for 12 percent, union variables 12 percent, taxes 53 percent, and employment protection for 24 percent of the increase in the unemployment rate. The long run multiplier shows that the increase in the benefit system accounts for 12 percent, unions 12 percent, taxes 63 percent, and employment protection legislation for 13 percent of the actual increase in the unemployment rate. Overall, the two methods for calculating long run effects give rather similar results. When comparing the results for

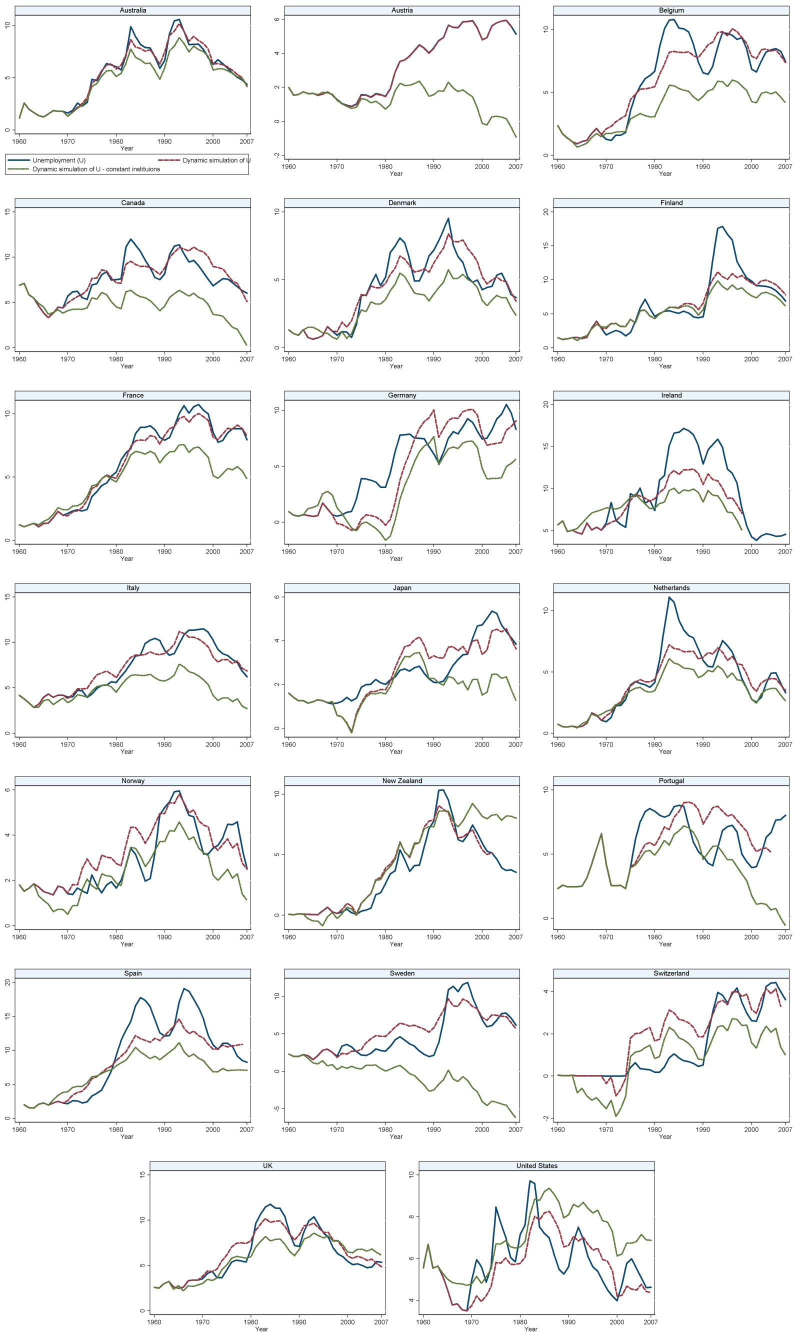

Figure 4: Dynamic simulation of equation (1) with constant and time-varying institutions on the revised and extended data set. Coefficients values from table 3 over the years 1960 to 2007. Percent

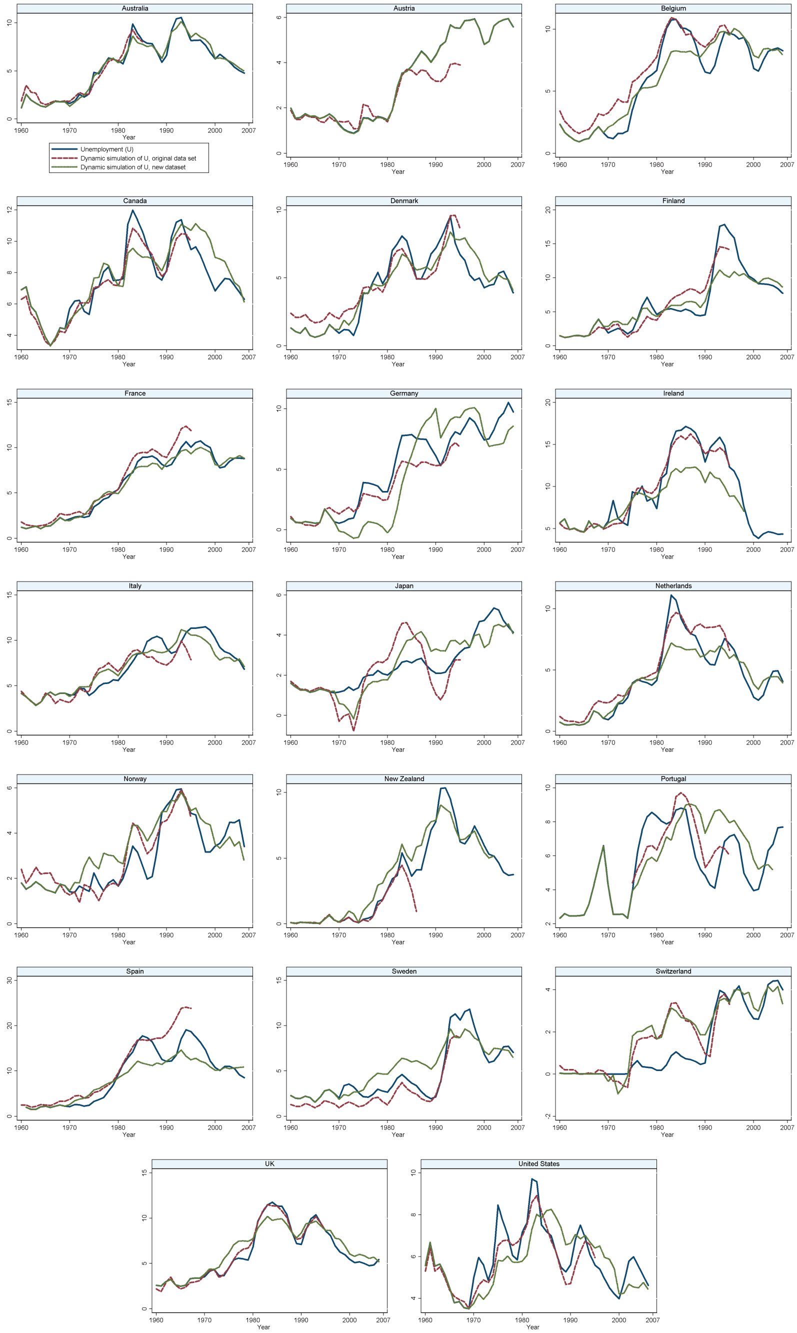

Figure 5: Dynamic simulation of equation (1) with time-varying institutions. Coefficient values from table 3 "1960-2007" and from the original estimation in appendix 2.C, model A in table C1. Percent

the 1960s to the average of 1990 to 1995 (table 2), the effects of benefits have decreased while the effect of taxes and employment protection has increased.

The analysis has not revealed the reason for the underprediction of unemployment that was found in section 2.4. However, a simulation of the unemployment rate seems to capture the development in actual unemployment also for the countries that has contributed to the decline in the unweighed average, Australia, Finland, France, Ireland, Italy, Netherlands, New Zealand, Spain and United Kingdom, exceptions might be Canada, Denmark, Norway and United States. The estimated effect of institutions still have the same direction of impact on unemployment. In fact, extending the sample suggests that some of the previous critique towards NNO is perhaps less serious when using the extended sample. For instance, Berger and Everaert (2008) claim that the cointegration of unemployment and institutions is spurious due to a non-rejection of a unit root in unemployment. However, the unweighed average of the unemployment rate has fallen since 1995, and a similar test on the extended data set would probably reject a unit root.

2.5 The dynamic specification of the NNO model.

To sum up so far: The underprediction is not due to the development in the included shock variables in NNO, as shown in appendix 2.B, figure B2. The data generating process has changed somewhat in the extended sample period, as shown by the estimated coefficient values displayed in table 3, and also the change to the original estimated coefficients as illustrated in figure 3. The new estimated coefficient values do not lead to a markedly better fit in the dynamic simulation over the extended sample period, as compared with the original coefficient estimates. The accounting exercise shows that changes in institutions can explain a large part of the actual change in unemployment similarly to the results of NNO. In fact, the analysis for the extended period shows that changing institutions account for a larger share of the increase in unemployment, but this is due to the actual increase being smaller. In this section, I explore another possible explanation for the underprediction, which is weaknesses in the dynamic specification. The first four sections 2.5.1 to 2.5.4 reveal weaknesses of the heterogenous specification of the model, while 2.5.5 shows that the model has unstable dynamics even if the heterogenous error term is neglected and the formal stability restrictions are met.

2.5.1 The solution of the full model

The empirical specification of NNO imposes the same coefficient values for all countries. Literally speaking, the interpretation is that the effect of, say, stricter employment protection on unemployment is the same in all countries. However, the specification of NNO also includes several country-specific terms, which leads to more flexibility across countries than one might expect given the assumption of common coefficients. The heterogeneity in the structural part of the model is included "... to ensure that the estimated coefficients are not distorted by omitted trended variables in each country or common shocks (NNO, page 15)". The heterogeneity specified in the error term is justified by a rejection of the joint hypothesis $\rho_i = 0$ for all countries in the panel (NNO, page 20). To illustrate the effect of the heterogenous lagged error term on the solution of the model for unemployment, equation (1) can be rewritten as follows:

$$U_{it} = \theta U_{i,t-1} + \boldsymbol{\beta X_{it}} + \boldsymbol{\alpha Y_{it}} + \gamma 1_t + \gamma 2_i + \gamma 3_i t + v_{it} \tag{3}$$

where $\boldsymbol{X}$ is a vector consisting of the institutional variables and the interaction terms in equation (1) and the vector $\boldsymbol{\beta}$ is the corresponding coefficients. The vector $\boldsymbol{Y}$ represents the shocks and the coefficients in vector $\boldsymbol{\alpha}$.

Second, by multiplying equation (3) with $(1 - \rho_i L)$ on both sides of this equation, using that $(1 - \rho_i L)v_{it} = \varepsilon_{it}$ defined by equation (2), we obtain

$$\begin{aligned}(1 - \rho_i L)U_{it} &= \theta(1 - \rho_i L)U_{i,t-1} + \boldsymbol{\beta(1 - \rho_i L)X_{it}} + \boldsymbol{\alpha(1 - \rho_i L)Y_{it}} \\ &\quad +(1 - \rho_i L)\gamma 1_t + (1 - \rho_i L)\gamma 2_i + (1 - \rho_i L)\gamma 3_i t + \epsilon_{it} \end{aligned} \tag{4}$$

Equation (4) has a white noise error term, ϵ_{it}. The equation can be organized as follows

$$\begin{aligned}(1 - (\rho_i + \theta)L + \rho_i \theta L^2)U_{it} &= \boldsymbol{\beta(1 - \rho_i L)X_{it}} + \boldsymbol{\alpha(1 - \rho_i L)Y_{it}} \\ &\quad +(1 - \rho_i L)\gamma 1_t + (1 - \rho_i L)\gamma 2_i + (1 - \rho_i L)\gamma 3_i t + \epsilon_{it} \end{aligned} \tag{5}$$

Equation (5) highlights important features of the specification used by NNO. The equation implies a 2nd order dynamic in the solution for the unemployment rate. This means that the stochastic process for the unemployment rate is an autoregressive moving average process with exogenous variables. This has some implications for the stability conditions and when using simulations to illustrate the fit of the model. It also implies that the estimated coefficients depend on some restrictions.

Equation (5) reveals that even if the coefficient vectors β and α by assumption are common across all countries, the dynamic effects of the explanatory variables will vary across countries. For the dynamic effects, the specification imposes the same relationship for all explanatory variables, but allows for variation among countries. To be concrete, equation (5) implies that an increase in, say, employment protection has the same direct effect on unemployment in all countries. However the effect of lagged employment protection on unemployment will vary across countries, depending on the country specific value of ρ_i. On the other hand, the relationship between short run and long run effects of the explanatory variables, eg. employment protection and the tax wedge, is assumed to be the same within each country, also given by ρ_i. This is not discussed in NNO.

2.5.2 Stability

The stability conditions of the homogenous part of the model depend on the coefficient value $|\theta| < 1$. This condition is met in NNO since $\theta = 0.86$. However, the stability conditions of both equation (1) and (2) can be divided into two parts; the stability of the homogenous part of the equation, i.e. equation (1), and the stability of the full model, i.e. equation (5). The second stability condition is not explored in NNO. It depends on the roots of the lag polynomial of order 2 being outside the unit root circle or equivalently that the corresponding characteristic polynomial lie inside the unit circle, see for instance Sydsaeter (1994, Ch. 6.5).

Table 5 lists the stability conditions in terms of ρ_i's to the solution of the estimated model in NNO. The conditions are calculated by finding stability conditions of the 2nd order differential equation in equation (5) with the estimated coefficients in appendix 2.C, in model A in table C1. All roots of the characteristic equation are real roots. If the values of the last three columns of table 5 are greater than zero, the 2nd order differential equation (5) will be stable. As seen from the table, Japan, New Zealand and Portugal have values lower than zero, implying that the stability conditions are not met. This means that the effect of the error term, for instance unspecified shocks, on unemployment increases over time. New Zealand has positive autocorrelation in the error term. This means that unemployment will increase or decrease steadily over time. Japan and Portugal have negative autocorrelation in the error term. This means that the error term in one period will increase unemployment but decrease unemployment in the next period. The positive and the negative effect of the error term to unemployment will increase over time.

Finally, note that an empirically unstable solution for unemployment is inconsistent with the equilibrium theory of unemployment. The effects are illustrated by a dynamic simulation in the next section.

2.5.3 Dynamic simulation rewritten

The simulations in NNO are based on equation (1), but equation (5) illustrates that the endogenous unemployment rate is also affected by the value of ρ_i. A simulation of equation (5) could give a more correct impression of how the model fits the data and how exogenous shocks not included in the empirical model will affect the unemployment rate in the long run.

Table 5: The error term and stability restrictions of equation (5) and the estimated coefficients in appendix 2.C, model A in table C1

							Stability	
Country	ρ_i	$a1$	$a2$	Rot 1	Rot 2	$1+a1+a2>0$	$1-a1+a2>0$	$1-a2>0$
Australia	0.39	-1.25	0.33	0.87	0.38	0.08	2.58	0.67
Austria	-0.19	-0.67	-0.16	0.86	-0.19	0.16	1.51	1.16
Belgium	-0.59	-0.27	-0.51	0.86	-0.59	0.22	0.76	1.51
Canada	-0.33	-0.53	-0.29	0.86	-0.33	0.18	1.24	1.29
Denmark	-0.49	-0.37	-0.42	0.86	-0.49	0.20	0.95	1.42
Finland	-0.98	0.11	-0.84	0.86	-0.97	0.27	0.05	1.84
France	0.11	-0.98	0.10	0.86	0.11	0.12	2.08	0.90
Germany	-0.34	-0.52	-0.30	0.86	-0.34	0.18	1.22	1.30
Ireland	-0.21	-0.65	-0.18	0.86	-0.21	0.17	1.46	1.18
Italy	-0.68	-0.19	-0.58	0.86	-0.67	0.23	0.61	1.58
Japan	-1.11	0.25	-0.96	0.86	-1.11	0.29	-0.21	1.96
Netherlands	-0.78	-0.08	-0.67	0.86	-0.78	0.25	0.41	1.67
Norway	-0.43	-0.43	-0.37	0.86	-0.43	0.20	1.06	1.37
New Zealand	2.92	-3.78	2.51	2.92	0.86	-0.27	7.29	-1.51
Portugal	-1.15	0.29	-0.99	0.86	-1.15	0.30	-0.28	1.99
Spain	-0.95	0.08	-0.81	0.86	-0.94	0.27	0.11	1.81
Sweden	-0.51	-0.35	-0.44	0.86	-0.51	0.21	0.91	1.44
Switzerland	-0.70	-0.16	-0.61	0.86	-0.70	0.23	0.55	1.61
United Kingdom	-0.41	-0.45	-0.36	0.86	-0.41	0.19	1.09	1.36
United States	-0.35	-0.52	-0.30	0.86	-0.34	0.18	1.22	1.30

A dynamic simulation of equation (5) with the estimated coefficient values in appendix 2.C, in model A table C1, is shown in figure 6 and in appendix 2.C, figure C3. The latter figure illustrates a similar pattern as the original simulations in appendix 2.C, figure C1 which were based on equation (1). This means that the effect of the heterogeneity in the error term on the simulations is small for a majority of countries. In contrast, for the countries with unstable solutions discussed in section 2.5.2 above, the effect of the heterogenous error term increases over time. As seen from figure 6, the simulated unemployment rate in New Zealand shows an explosive path. For Japan and Portugal there is a negative autocorrelation, and for Portugal the positive and negative effects increase over time.

The simulations illustrate that the empirical model has an unstable solution for three of the countries when the error term is seen as an integrated part of the econometric model as in equation (5). The simulations illustrates the instability results for these countries in table 5.

2.5.4 The heterogeneity and estimation method

As mentioned in the beginning of this section, the coefficients of both the lagged endogenous variable and the exogenous variables in equation (5) implicitly depend on the relationship between the current and lagged values of the variables included in the model. The relationship between the current and lagged value of the included variables is the same for all variables captured by the common factor $(1-\rho_i L)$. One should note that the

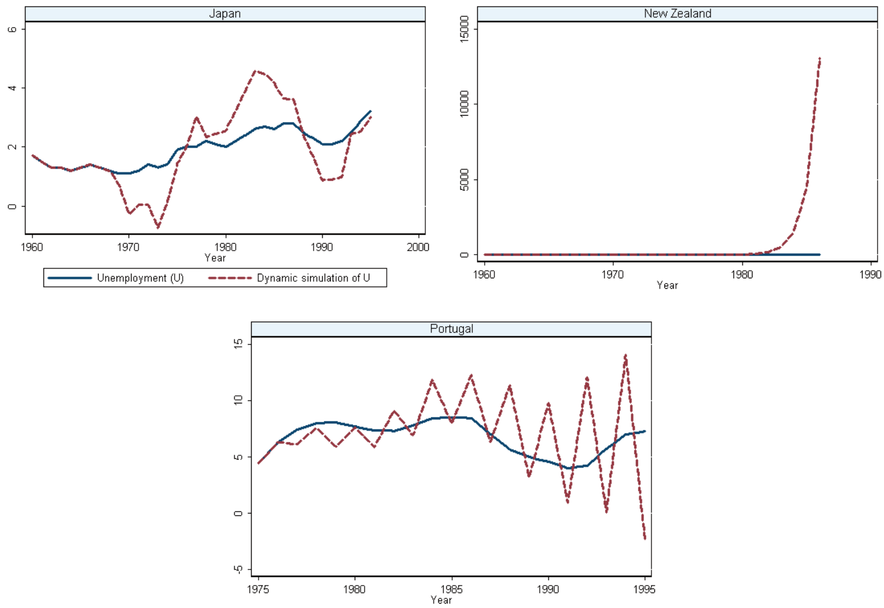

Figure 6: Simulations of equation (5). Estimated coefficients in appendix 2.C, in model A in table C1. Percent

feasible general least squares with fixed effects which is used to estimate the coefficients, normally results in a bias for the lagged endogenous variable. The bias occurs since the value of the unemployment rate in the previous period per assumption is correlated with the error term, see for instance Baltagi (2008, Ch. 8). Since both θ and ρ_i are estimated jointly, a bias in the value of θ could possibly affect the value of ρ_i and hence the value of the coefficients of the exogenous variables.

An equivalent procedure to the feasible general least squares estimation method used by NNO, is to perform a within-transformation of equation (3), (Baltagi (2008, Ch. 3)):

$$\begin{aligned} U_{it} - \overline{L.U}_{i.} &= \theta(U_{i,t-1} - \overline{L.U}_{i.}) + \boldsymbol{\beta}(\boldsymbol{X_{it}} - \overline{\boldsymbol{L.X}}_{\boldsymbol{i.}}) + \boldsymbol{\alpha}(\boldsymbol{Y_{it}} - \overline{\boldsymbol{L.Y}}_{\boldsymbol{i.}}) \\ &\quad +(\gamma 1_t - \overline{\gamma 1}_{.}) + \gamma 3_i(t - \bar{t}) + (v_{it} - \overline{L.v}_{i.}) \qquad (6) \end{aligned}$$

$U_{i,t-1} - \overline{L.U}_{i.}$ is the unemployment rate in the previous period minus the average sum of unemployment summed over all time periods up to period $t-2$. This term is correlated with $v_{it} - \overline{L.v}_{i.}$ by definition. Since θ is positive, this implies that the estimated persistence will be underestimated. Since the error term is autocorrelated in the empirical specification and estimated jointly with the autoregressive parameter, the problem is even more severe given the difficulty of deriving a consistent estimate of the AR parameters in that context.

If the estimated value of θ are biased, the ρ_i's are probably biased too. Due to the common factor described above, the estimated coefficients of the explanatory variables might also be affected. In addition, if some of the regressors are correlated with the lagged dependent variable to some degree, their coefficients may be seriously biased for this reason as well.

One solution to the bias problem that arises with the combination of an endogenous lagged variable and country specific shifts, is to take 1st order differences of the original model, see Baltagi (2008, Ch. 8). The 1st difference transformation removes the country specific effect. There is still a correlation between the difference lagged dependent variable and the disturbance process (which now is a 1st order moving average, MA(1)), but now an instrument variable is available. This is true even if the v_{it} follows an AR(1) process. In the model described in NNO the error term follows an AR(1) process.

$$\begin{aligned} U_{it} - U_{it-1} &= \theta(U_{i,t-1} - U_{it-2}) + \boldsymbol{\beta}(\boldsymbol{X_{it}} - \boldsymbol{X_{it-1}}) + \boldsymbol{\alpha}(\boldsymbol{Y_{it}} - \boldsymbol{Y_{it-1}}) \\ &\quad +(\gamma 1_t - \gamma 1_{t-1}) + \gamma 3_i(t - (t-1)) + (v_{it} - v_{it-1}) \end{aligned} \tag{7}$$

since $(\gamma 1_t - \gamma 1_{t-1}) = \gamma 1_t$ and $t - (t-1)) = 1$ the equation can be reduced to

$$\begin{aligned} U_{it} - U_{it-1} &= \theta(U_{i,t-1} - U_{it-2}) + \boldsymbol{\beta}(\boldsymbol{X_{it}} - \boldsymbol{X_{it-1}}) + \boldsymbol{\alpha}(\boldsymbol{Y_{it}} - \boldsymbol{Y_{it-1}}) \\ &\quad +\gamma 1_t + \gamma 3_i + (v_{it} - v_{it-1}) \end{aligned} \tag{8}$$

One way to use the 1st difference approach on the specified model in section 2.2 is to assume a second order dynamic equation (according to the dynamics in the rewritten equation (5)) with or without heterogenous error terms.

However, one should be aware of that the bias caused by the endogenous lagged variable and the fixed effect decreases with a increase in the sample period, see Judson and Owen (1999). The fixed effect bias on OECD panel data is also discussed in Nymoen and Sparrman (2010).

2.5.5 The dynamics of the homogenous part of the empirical model in NNO

The difference between the simulated and the actual value apparent in figure 2 of unemployment is still not found. It is therefore interesting to check if some of the difference could be explained by the dynamics of the specified model in NNO.

Figure 7 shows the simulated unemployment rate out of sample for the dynamic part of the equation, which means that only the lagged unemployment rate, the trend and the fixed effects are included in the simulation. The figure illustrates that the unemployment

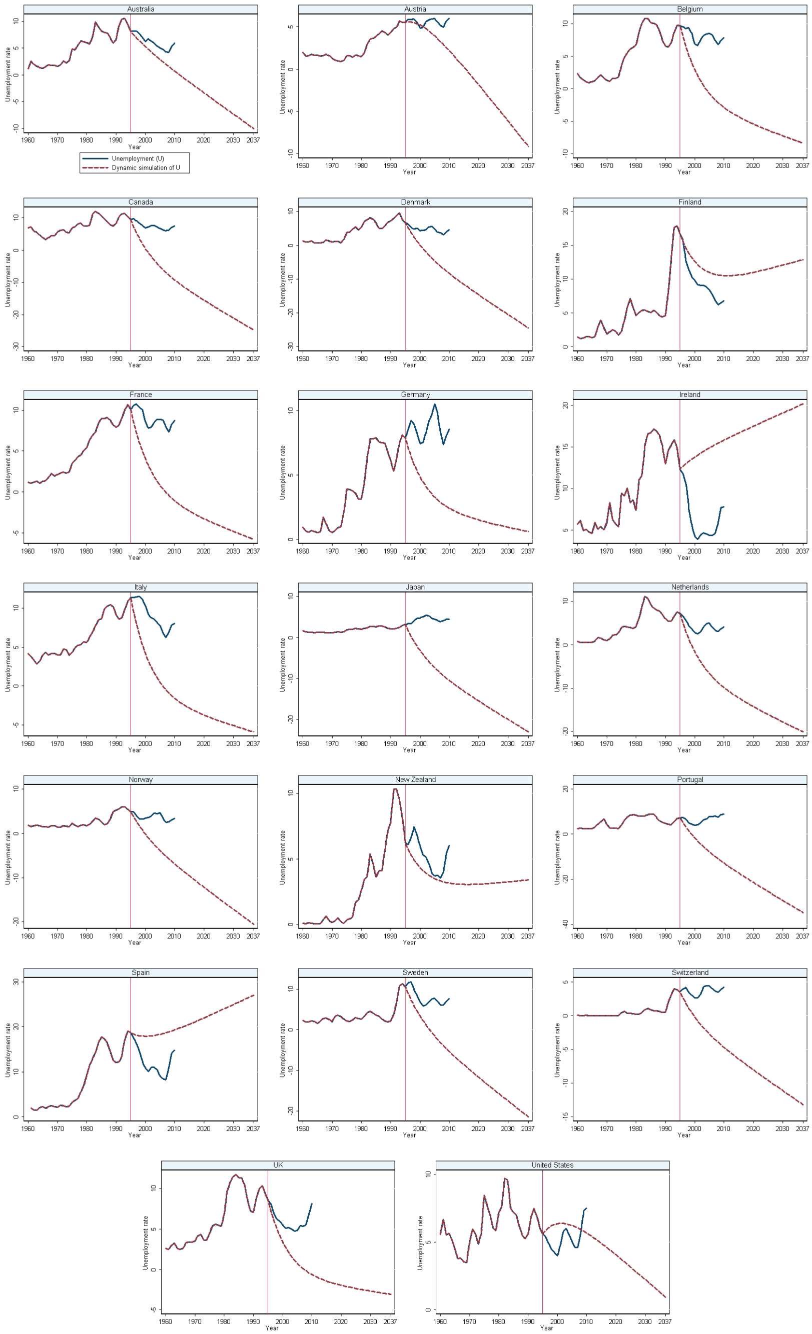

Figure 7: Dynamic simulation of $\tilde{U}_{it} = \hat{\theta}\tilde{U}_{i,t-1} + \hat{\gamma 2}_i + \hat{\gamma 3}_i t$ on the revised and extended data set. Coefficient estimates from appendix 2.C, model A in table C1, over the years 1960 to 2037. Percent

rate out of sample is largely driven by this dynamics, where only New Zealand seems to have a stable path for the rate of unemployment. Thus, the large underprediction seems to be the result of unstable dynamics. Stable dynamics is essential for the ability to predict a stationary time series such as unemployment by simulation.

2.6 The effect of changes in labour market institutions

In view of the strong and in some cases erratic effects of the dynamic specification of NNO, it is useful to explore the effect of the institutional variables on the rate of unemployment by other methods. In this section, I explore the predicted change in the rate of unemployment from the NNO model, taking into account only the changes in the institutional variables. More specifically, for each country I compute the change in unemployment that follow from the change in the institutional variables from 1995 to the 5 year average in the period 2002 to 2007 (2007s hereafter), as measured by long run multiplier using the coefficient estimates in appendix 2.C, in model A in table C1. Table 6 presents the results for all countries put together, but viewing one institutional variable at the time. We see from line 1 that on average, employment protection legislation has become somewhat less strict, with a reduction in the average index value of 0.10. Given the estimated coefficients, this should lead to a long run reduction of the rate of unemployment of 0.11 percentage points. The long run effect of the change in all institutional variables from 1995 to 2007s is an increase in unemployment of 1.28 percentage points, cf. the last row in table 6. In contrast, actual unemployment fell by 2.3 percentage points over the same period.

At first thought, the difference between the predicted increase and actual fall in unemployment does not look good for the predictive power of the NNO model. However, by matching separate computations for each country, calculating the predicted change in unemployment due to changes in all the institutional variables for the country, a more flattering picture emerges. Figure 8 displays a cross plot of the predicted change in unemployment due to institutional changes and the actual change in unemployment, for all European countries and for the whole sample. The positive relationship is clear, and the slope is estimated to be 0.4; is significant at a 5 percent level if Ireland is treated as an outlier (significant at a 10 percent level if the slope is estimated on the whole sample). This is a fairly strong positive correlation, and 9 out of 14 countries are in the predicted quadrants (ie. upper right or lower left). In particular Denmark and Finland have changed their labour market institutions in a way that should give lower unemployment, and these countries also experienced a large reduction in actual unemployment. In contrast, Germany and Portugal have changed their institutions in the opposite direction, and unemployment has increased. The positive correlation suggest that changes in institutions do explain an important part of the changes in unemployment. However, as compared to an upward sloping 45° degree line through origin, which would prevail in the hypothetical case where the model were correct and no other changes took place, we also observe three discrepancies. First, the observations are too low, indicating that unemployment has fallen for other reasons. A likely reason is the fairly positive overall evolution of most OECD economies over this period, with considerable growth of GDP. Second, we observe two notable outliers, Ireland and Spain, which have experienced a large reduction in unemployment in spite of the adverse change in labour market institutions. Given the subsequent sharp deterioration of the Spanish and Irish economies, it seems reasonable to explain part of the reduction in unemployment as caused by unsustainable overheating of the economy. Adjusting for this would make these countries more in line with the other countries. Third, we observe that the slope is less steep than 45°. However, one should remember that while the figure is based on long run effects, the full impact of institutional changes late in the sample period will not be reaped in 2007, implying that the figure in fact may exaggerate somewhat the predicted effect of the institutional variables. Note also that the scale of the induced change in unemployment, calculated by the long run multiplier, depends crucially on the estimated value of the lagged unemployment rate coefficient. As discussed above, this value could be biased, and vary between countries if the heterogeneity is taken into account. This means that the total effect of institutions, and general validity of institutions effects on unemployment should be interpreted with care.

Table 6: The predicted effect on unemployment calculated using actual changes in institutions from 1995 to 2002-07 and the estimated coefficients in appendix 2.C, model A in table C1

Institutional variable X:	Coefficient values from table C1 in appendix 2.C	Actual change in institution X; $\Delta X = \bar{X}^a_{07-02} - X_{95}$	Contribution of institutions to U The long run multiplier[b]: percentage points	share of total[c]
Employment protection (EPL)	0.15	-0.10	-0.11	-0.09
Benefit replacement ratio (BRR)	2.21	0.02	0.36	0.28
Benefit duration (BD)	0.47	0.12	0.41	0.32
Interaction BRR and BD	3.75	0.00	0.08	0.06
Interaction CO and $UDNET$	-6.98	-0.01	0.56	0.44
Interaction CO and TW	-3.46	0.01	-0.37	-0.29
Union density, 1st. diff. ($\Delta UDNET$)	6.99	0.01	0.28	0.22
Coordination (CO)	-1.01	0.04	-0.30	-0.23
Tax rates (TW)	1.51	0.03	0.37	0.29
Sum			1.28	1.00

a) $\bar{X}_{07-02}$ is the average level of the institutional variable in the period 2002 to 2007.
b) The long run multiplier is calculated by the use of formula;
$\frac{\hat{\beta} * \Delta X}{1-\theta}$ with the actual change in the specific institution (X).
c) Share of total increase in unemployment rate induced by changes in institutions.

Table 7 shows the decomposition of the predicted effect on unemployment to specific

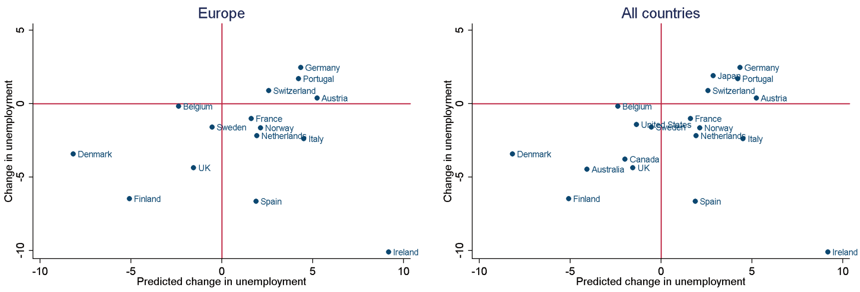

Figure 8: The predicted change in unemployment is calculated by the long run multiplier, using the formula; $\frac{\hat{\beta}*\Delta X}{1-\theta}$ with the actual change in the specific institution (X) from 1995 to 2002-07. The changes in institutions are calculated on the revised and extended data set, and β and θ are the estimated coefficients in appendix 2.C, in model A in table C1

Table 7: The predicted effect on unemployment by each institutional variable calculated using the actual changes in institutions from 1995 to 2002-07 and the estimated coefficients in appendix 2.C, in model A in table C1

Country	**Empl. prot.**	**Benefit repl. (BRR)**	**Benefit duration (BD)**	BRR **and** BD	CO **and** $UDNET$	CO **and** TW	**Union density ($\Delta UDNET$)**	**Coord. (CO)**	**Tax rate** TW	**Tot**
Austria	-0.1	1.4	0.1	0.7	3.6	-0.3	-0.3	0	0.1	5.2
Belgium	-0.4	-0.3	0.1	-0.1	-0.7	-0.3	-0.3	-0.7	0.2	-2.4
Denmark	0	-0.5	-0.3	-1.2	-3.2	-1.3	-0.2	-1.8	0.4	-8.2
Finland	-0.1	-5.1	0.9	-1.3	0.7	-0.1	-0.4	0	0.1	-5.1
France	0	0.6	0.4	0.7	0	0	0	0	0	1.6
Germany	-0.3	0.3	0.7	-0.1	3.6	-0.2	0.2	0	0.1	4.3
Ireland	0.1	0.6	0	0.4	5	-1.6	4	0	0.7	9.2
Italy	-0.6	8.1	1.4	-3.2	2	-2.2	0.3	-2	0.9	4.5
Netherlands	-0.2	0.2	0.6	1.4	-0.2	0	-0.2	0	0.3	1.9
Norway	0	0.5	0.3	0.6	0.6	-0.2	0.2	0	0.2	2.1
Portugal	-0.1	0.8	0.9	2.1	-0.2	0.1	-0.2	0	0.7	4.2
Spain	0	-0.2	0.2	0.5	0	0.1	0.5	0	0.8	1.9
Sweden	-0.1	-1.1	0	0.7	-0.3	0.1	-0.4	0	0.6	-0.5
Switzerland	0	0.7	0.5	0.9	0.3	-0.1	0	0	0.2	2.6
UK	0	-0.5	0.4	-0.9	-2.2	0.4	1	0	0.2	-1.5
Total	-0.1	0.4	0.4	0.1	0.6	-0.4	0.3	-0.3	0.4	1.3

institutions for the European countries in the sample. Increases in the benefit system for the unemployed workers and increases in unions density are the largest contributors to the predicted increase in unemployment. Reduction in employment protection and increase in coordination among the wage setters have lowered the predicted increase in unemployment. The table also reveals the large variation in the institutional development among the countries in the sample.

2.7 Conclusion

This paper has replicated and assessed the model of Nickell et al. (2005). Specifically, I use the replicated model to predict unemployment out of sample, benefitting from twelve additional years of data. The dynamic simulation reveals that the model in Nickell et al. (2005) is not useful to forecast the evolution of the unemployment rate in the post-estimation period as it delivers a severe underprediction of the change in unemployment.

For most countries, the underprection is reduced, although still quite large, when taking the change in institutions into account. This is the case for Australia, Austria, Belgium, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, Norway, Portugal, Sweden, and Switzerland. For Ireland, there is no difference between simulated unemployment rate with constant and time-varying institutions. For other countries, Finland, New Zealand, Spain, the United Kingdom and the United States, the fit to the actual unemployment rate worsens when taking the change in institutions into account.

I then explore possible explanations for the underprediction. One obvious candidate is that the shocks that are included as explanatory variables in the empirical model have evolved differently in the post sample period and caused underprediction of the unemployment rate. However, the results of a simulation with variation in the shock variable and a simulation where the shock variables are set to zero in the post sample period, shows nearly no effect of the shocks in the extended period, with Japan and Italy being the only exceptions.

Another possible explanation for the underprediction could be that the the data generating process has changed over time. This might be in line with some of the earlier critique, which have argued that there have been limited possibilities to detect plausible effects of institutions to unemployment due to short time series (Belot and van Ours, 2004). In order to investigate this possible explanation, the model in Nickell et al. (2005) was reestimated on the revised and extended data set. The reestimation showed that the size of the coefficients changed quite substantially both on the revised data set and on the extended time period. However, dynamic simulation does not give a better fit with the reestimated coefficients than with original ones, even for the extended sample period, suggesting that changes in the coefficient values is not the key explanation for the underprediction.

The model dynamics have been investigated to detect if this is the cause of the underprediction in the post sample period on the original data set. A dynamic simulation of the full model, when the error term is explicitly taken into account, reveals that the model is non stationary for some countries. This is verified by the roots of the 2nd order differential equation that is implied by the model in Nickell et al. (2005). The results suggest a reconsideration of the dynamic specification of the model. Especially since the underlying solution to the specified model implies a 2nd order dynamics in the unemployment rate. However, this is not the main cause of the underprediction of the unemployment rate, as most of the countries have a stable solution of the model. Instead it turns out that the underprediction of the unemployment rates is largely driven by the dynamic specification of the model, where the combination of the large coefficient for the lagged unemployment rate, the trend and the fixed effects, implies a tendency for unemployment to diverge in one direction or the other. This implies that forecasting a stationary time series as the unemployment rate is impossible.

One could argue that the model in Nickell et al. (2005) was developed to explore the link between institutions and unemployment and not to predict unemployment. What about the link between institutions and unemployment, which was the main topic of interest for Nickell et al. (2005)? Repeating their analysis of the long run effect of changes institutions for the extended sample period for the European countries, I find that changes in institutions now account for 76 percent of the total change in unemployment from the 1960s to 2002-2007, up from 41 percent in my replication of their results over the shorter time period until 1990-1995. At face value, this might suggest that institutions have become more important. However, the interpretation is less clear cut. First, the larger share reflects that unemployment increased less over the longer period, so there is less to explain. More importantly, looking only at the effect of the change in institutions, this would actually predict that unemployment increased by 1.3 percentage points over the period 1995 to 2002-2007, while actual unemployment instead fell by 2.3 percentage points. Thus, the sign in this aggregate relationship is wrong.

However, taking into account country variation, a different picture emerges. There is a clear tendency that countries which have changed their institutions in an “employment-friendly” way, like Denmark and Finland, have experienced a larger reduction (or smaller increase) in unemployment than the countries that have changed their institutions in the opposite direction, like Germany and Portugal. This is a clear indication that labour market institutions affect unemployment in the direction found by Nickell et al. (2005), even if the large underprediction of unemployment for the majority of the countries shows that their model is unable to account for the overall evolution of the rate of unemployment.

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2.A The Data: Definitions and sources

This appendix contains information about variables that are important for the evolution of the unemployment rate in 20 OECD countries. The countries in the sample are:

Australia	Finland	Japan	Spain
Austria	France	Netherlands	Sweden
Belgium	Germany	Norway	Switzerland
Canada	Ireland	New Zealand	United Kingdom
Denmark	Italy	Portugal	United States

The variables are collected in two data bases. The first data base contains the original data set from Nickell et al. (2005) and contains observations from 1960 to 1995. The definitions are as given in the appendix in Nickell et al. (2005). The second data base, referred to as the revised and extended data set hereafter, contains revised and extended time series for the same variables as the original data set. The sources of the second data base is OECD (2002), OECD (2006), OECD (2008b), OECD (2008a) and Nickell (2006).

The revised and extended data set follows the set up in Nickell et al. (2005) where the variables are divided into two groups; economic variables described in section 2.A.1 and institutional variables described in section 2.A.2. The extended at revised data set contains observations from 1960 to 2007.

2.A.1 Economic variables

The economic variables are available at a yearly frequency in OECD (2008a)[5] and missing observations are replaced with observations from earlier data bases OECD (2002), OECD (2006) and OECD (2008b).[6]

U: Unemployment rate

The standardized unemployment rate (UNR) in Economic Outlook OECD (2008a) is used as a primary data source for the unemployment rate in the OECD countries, and missing observations are replaced by the growth rate in a corresponding time series in an earlier data base, OECD (2002). Australia, Denmark, Germany, Spain and Switzerland are prolonged by the formula in equation (A1):

$$Y_{it} = Y_{it+1} * \frac{X_{it}}{X_{it+1}} \tag{A1}$$

where Y_{it} denotes (UNR) in OECD (2008a) and X_{it} denotes the (UNR) in the earlier data base OECD (2002) for country i in time period t.

[5]Data are collected and organized by the author. This implies that neither OECD nor any other source is responsible for the analysis or the interpretation of the data in this paper.

[6]An comprehensive overview of data and data sources are available upon request.

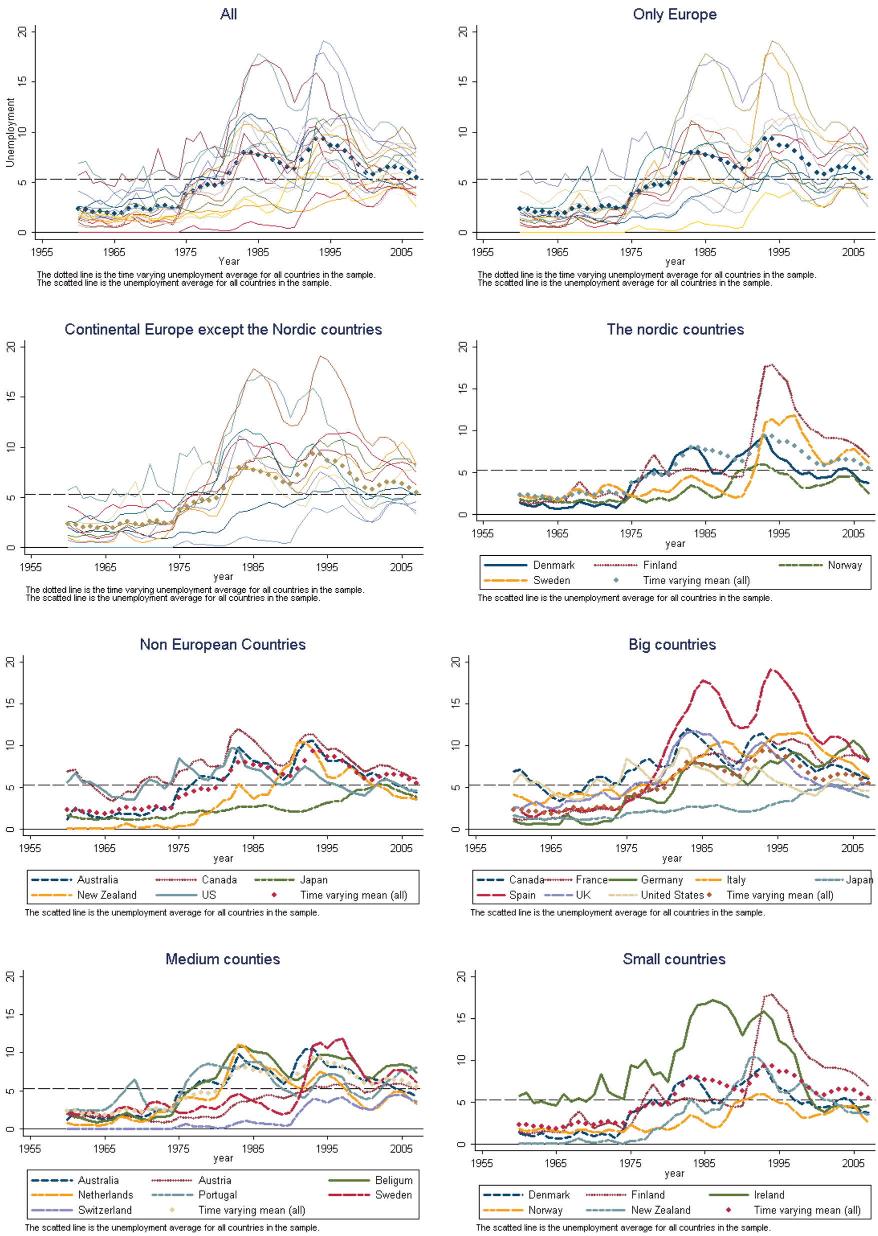

Figure A1: Actual unemployment. Revised and extended data set over the years 1960 to 2007. Percent

Australia and Denmark are prolonged five years backwards. Germany from 1991, Spain from 1976, Switzerland from 1969 and backwards.

The revised and extended data set has one less observation for Spain in 1960 than the data base from Nickell et al. (2005).

RIRL: Real interest rate

The real interest rate is calculated in two steps from the nominal interest rate and the consumer price index. First, inflation is calculated by the change in the time series for consumer price index as in equation (A2). Then, the real interest rate is defined as the nominal interest rate minus the inflation rate as in equation (A3).

$$\pi_{it} = 100 * \frac{\Delta CPI_{it}}{CPI_{it-1}} \tag{A2}$$

Inflation for country i in time period t, π_{it}, is the difference between the price level this period, CPI_{it}, minus the price level previous period, CPI_{it-1}, divided by the price level previous period. The real interest rate, r_{it}, is then calculate by subtracting the inflation rate from the long term nominal interest rate, IRL_{it},.

$$r_{it} = \frac{IRL_{it} - \pi_{it}}{100} \tag{A3}$$

The main source for the long term nominal interest rate on government bonds, IRL, and the consumer price index, CPI is OECD (2008a). CPI for United Kingdom is from OECD (2006) with the last observation is in 2008. CPI for Spain is from OECD (2007) with the last observation is in 2009. For Germany before 1991, CPI is prolonged with the growth rate in the same time series in data base OECD (2002), by equation (A1) where Y is CPI in OECD (2008a) and X is CPI in OECD (2002). Note that the base year in CPI varies between countries and data bases. The base year is however not important since we are only interested in the price growth.

The missing observations in the nominal interest rate for Germany are replaced by the growth rate in IRL in OECD (2002) by equation (A1) where Y is IRL in OECD (2008a) and X is CPI in OECD (2002).

Compared with the real interest rate in the original data set in Nickell et al. (2005): 1 observation in 1969 is missing for Australia. One observation in 1960 for France, Germany, Italy, Sweden, Switzerland, UK and United States is missing. The years 1960 to 1970 are missing for Ireland. Observations for the years 1960 to 1961 are missing for Netherlands. Finally the years 1961 to 1969 are missing for New Zealand.

$PROD\_hp$**: Productivity shock**

Productivity shocks are measured by the difference between actual productivity and the productivity trend. Productivity trend is calculated by a Hodrick Prescott filter of log real gross domestic product, minus log of total employment. The trend is defined by equation (A4):

$$PROD\_hp_t = HPtrend[log(GDPV_t) - log(ET_t)] \tag{A4}$$

where $GDPV$ is real gross domestic product and ET total employment.

The main source for both variables the real gross domestic product and total employment are $GDPV$ and ET in OECD (2008a). Missing observations in Denmark, Germany and Switzerland for $GDPV$ are replaced by observations in OECD (2002) adjusted by differences in mean and slope in the time series as in equation (A5).

$$Y1_{it} = Y1_{it+1} * \frac{(X1_{it} * b_i + a_i)}{(X1_{it+1} * b_i + a_i)} \tag{A5}$$

where $Y1$ is $GDPV$ in OECD (2008a), $X1$ is $GDPV$ in OECD (2002), and a_i and b_i are the estimated coefficients from a regression between the two time series for country i. The regression reveals significant differences in both coefficients for most countries. a_i is the average difference in level and b_i the average difference in slope between the two series for country i.

The missing observations for ET in Australia before 1964, Denmark before 1966, Germany before 1991 and Switzerland before 1970 are replaced by the formula in equation (A1) where Y is ET from OECD (2008a) and X is ET from in OECD (2002).

LDS: Labour demand shock

Labour demand shocks are defined as the residuals, $\hat{\epsilon}_t$, from the following regression:

$$\begin{aligned} log(ET_{it}) = \beta0 + \beta1 log(ET_{it-1}) + \beta2 log(ET_{it-2}) + \beta3 log(ET_{it-3}) \\ + \beta4 log(GDPV_{it}) + \beta5 log(w_{it}) + \epsilon_{it} \end{aligned} \tag{A6}$$

where ET is total employment and $GDPV$ is real gross domestic product as the variables defined under the productivity shock above. w is labour cost and defined by equation (A7).

$$w = log(IE) - log(ET) - log(PGDP) \tag{A7}$$

where, IE is compensations of employees, ET is total employment and $PGDP$ is gross domestic product deflator at market prices.

The data source of the variable ET are described in the previous section, i.e. productivity shocks.

The main data source for compensations of employees is IE in OECD (2008a), but the main source is IE in OECD (2006) for New Zealand and contains observations from 1986 to 2008. Compensations of employees is prolonged in Austria before 1964, Germany before 1991, Norway in the period 1962 to 1974, New Zealand in the period 1971 to 1986 and Switzerland before 1990 by the formula in equation (A5), where $Y1$ is IE in OECD (2008a) and $X1$ is IE in OECD (2002).

The main data source for $PGDP$ is OECD (2008a). The deflator is prolonged backward in Denmark before 1966, France before 1993, Germany before 1991 and Switzerland before 1966 by equation (A1) where Y is the $PGDP$ in OECD (2008a) and X is the $PGDP$ in OECD (2002) when observations are missing.

Note also that the deflator used here is different from the GDP deflator at factor costs used in the original data set. It could be argued that this deflator is more consistent, since the variable IE includes the effect of value added tax.

Compared with the original data set, there are still some observations missing for the labour demand shock. The missing observations are replaced by the shocks in Nickell et al. (2005) for Canada in 1960, Denmark in 1960 to 1965, France in 1960 to 1962, Norway 1960 to 1962, New Zealand 1960 to 1971 and Portugal in the period 1960 to 1982.

d2MS: Acceleration in money supply

Acceleration in money supply is equal to the second difference of the money supply $\Delta 2MS$.

The main source for money supply is $MONEYS$ in OECD (2008a). Observations for $MONEYS$ in Australia, Canada, Germany, New Zealand, Uk and Sweden are from OECD (2008b). Time series for $MONEYS$ in Denmark, Spain and Switzerland are from OECD (2006). $MONEYS$ from OECD (2008a) and OECD (2008b) are first prolonged backwards with the growth rate in $MONEYS$ in OECD (2006) when differences in mean and slope is adjusted for as in equation (A5). Then, the prolonged $MONEYS$ from OECD (2008a), OECD (2008b) and OECD (2006) are prolonged backwards with the growth rate in $MONEYS$ in OECD (2002) when differences in mean and slope is adjusted for as in equation (A5).

Still, there are some missing observations. The missing observations are therefore replaced backwards with the shocks in the original data set: Belgium in the period 1962 to 1970, Canada in the period 1962 to 1969, Denmark in the period 1962 to 1965, Finland in the period 1962 to 97, France in the period 1962 to 1978, Germany in the period 1962 to 1970, Italy in the period 1962 to 1976, New Zealand in the period 1962 to 1966, Spain in the period 1962 to 1964 and United Kingdom in the period 1962 to 1964.

TTS: Terms of trade shock

Terms of trade shock is calculated by the change in the import price deflator relative to the real GDP deflator at market prices multiplied with the share of imports to GDP. The following equation (A8) calculates the shocks:

$$TTS = \frac{MGS}{GDP}\Delta\{ln(\frac{PMGS}{PGDP})\} \tag{A8}$$

MGS is imports at current prices, GDP is GDP at current prices, $PMGS$ is import price deflator and $PGDP$ is the real GDP deflator at market prices described under labour demand shock.

The sources of the time series for $PGDP$ are previously described under labour demand shock. The main source for import price deflator and GDP at current prices is $PMGS$ and GDP in OECD (2008a). GDP and $PMGS$ in OECD (2002) is used to prolong the variables from OECD (2008a) backwards when observations are missing by equation (A5).

The replaced observations for GDP and $PMGS$ are Denmark before 1966, Germany before 1991 and Switzerland before 1965. MGS in OECD (2002) prolongs the time series for Denmark before 1966, Germany before 1991 and Switzerland before 1965 by equation (A5).

2.A.2 Structural variables

New information for institutional variables is available every second or fifth year. The main data source is OECD (2004), but due to missing observation some data are collected from Nickell (2006). The time series are replaced by data from other sources when observations are missing. The variables and the method for prolonging are discussed in detail in the next sections.

TW: Tax wedge

The rates described here are calculated from actual tax payments. The total tax wedge is equal to the sum of the employment tax rate (t1), the direct tax rate (t2) and the indirect tax rate (t3), as given in equation (A9).

$$TW = t1 + t2 + t3 \tag{A9}$$

$t1$ is equal to employers total wage costs calculated by the sum of wages received by employees and taxes payed by the employer to the government. This gives the following relationship; $t1 = SSRG/(IE - SSRG)$, where SSRG is social security contributions and IE is compensation to employees. The latter consist of two main components, wages and salaries and social contributions. Social contributions are payed by the employers to social security schemes or private funded social insurance schemes. $t2$ are direct taxes payed by the households (TAXh) divided by current receipts of households (CRh), i.e. $t2 = TAXh/CRh$. Finally $t3 = (TAXind - SUB)/Cp$, where TAXind are net indirect taxes, SUB is the value of subsidies and Cp is the value of private final consumption expenditure.

The main data source for tax wedges is OECD (2008c) which contains information for the period 1960 to 2010. The latter years are predictions. The tax rates are calculated by the above formulas, and when a tax rate is missing, the growth rate in the same tax rate but from the data base of Nickell (2006) in the period 1960 to 2003 is used to prolong the time series for the following countries: Belgium is prolonged before 1965, Denmark is prolonged before 1966, Germany before 1970, Portugal is prolonged in the period 1960 to

1995 and Switzerland is prolonged before 1990 with the tax rates in OECD (2008c). Tax rates for Australia, Austria, Canada, Finland, France, Ireland, Italy, Japan, Netherlands, Norway, Spain, Sweden, the United Kingdom and the United States are not prolonged and are taken directly from the main data source OECD (2008c). New Zealand has the main data source Nickell et al. (2005) due to missing observations, and before 1975 the growth rate in the sum of $t1$ and $t2$ from the source Nickell (2006) are used to prolonged the time series backwards. The tax series for New Zealand is then prolonged forward with the growth rate in $t3$ from source Nickell (2006) after 1986. Note also that the $t3$ is interpolated due to one missing observation in 1991.

BRR: Benefit replacement rates

The benefit replacement ratio is a measure of how much each unemployed worker receives in benefits from the government. The benefit replacement ratio is described in detail below.

The detailed rate for unemployment benefits divides data in three different family types: single, with a dependent spouse and with a working spouse. The benefits also depend on the employment situation: 67 percent and 100 percent of the average earnings. Within these groups, benefits are divided into the duration of benefits when being unemployed. One variable for how much each of the former groups receives in the first year, the second and third year and the fourth and fifth year. The indexes used in this paper uses the indexes aggregated over over family types. This results in six different groups: brr67a1, brr67a2, brr67a4, brr100a1, brr100a2 and brr100a4.

brr67a1: First year benefit replacement rate for workers with 67 percent of average earnings and the average over family types.

brr67a2: Benefit replacement for the second and third year. 67 percent of average earnings and the average over family types.

brr67a4: Benefit replacement rate for the fourth and fifth year. 67 percent of average earnings and the average over family types.

brr100a1, brr100a2 and brr100a4: The same as the former but for 100 percent of average earnings.

The average of brr67a1 and brr100a1 is used as an indicator of benefit replacement ratios, i.e. BRR.

The main source for the more detailed benefit ratios is tables in employment outlook, see OECD (2004). Observations are provided every second year from 1961 to 2001. The time series are interpolated over the years, and extracted by the last known observation.

BD: Benefit duration

Benefit duration is a measure of how long the benefits last when being unemployed. The ratio is calculated by the time series described under benefit replacement ratio, by equation (A10).

$$BDj_{it} = \alpha \frac{brrja2_{it}}{brrja1_{it}} + (1-\alpha)\frac{brrja4_{it}}{brrja1_{it}} \tag{A10}$$

where $\alpha = 0.6$, $j = \{67, 100\}$, $i = 1, 2...20$ and $t = 1960, 1961...2007$. $brrja1_{it}$ is the benefit replacement rate in year 1, $brrja2_{it}$ is the benefit replacement rate in year 2 and 3, and finally $brrja4_{it}$ is benefit replacement rate in year 4 and 5. $\alpha = 0.6$ gives more weight to the second and third year as compared to the fourth and fifth year. The index is calculated for both employment situations, i.e. $j = 67$ percent and $j = 100$ percent of average earnings.

If benefit duration stops after one year, then brr67a2 = brr67a4 = 0, and BD67 = 0. If benefit provision is constant over the years, then brr67a1 = brr67a2 = brr67a4, and BD67 = 1. However, some countries increase payments over time and the value of benefit duration is above one.The average of $bd67_{it}$ and $bd100_{it}$ is used as an indicator of benefit duration, i.e. BD_{it}.

UDNET: Union density

Union density rates are mainly constructed using the number of Union memberships divided by the number of employed, see Visser (2009).

The database Nickell (2006) contains additional information for Sweden before 1975 and Ireland in 1960. The time series for Sweden in the latter source is first interpolated and the growth rate is calculated to prolong the original time series from Visser (2009).

It is the first difference in union density that is used in the regressions in the paper, $\Delta UDNET$.

By comparing the time series with the original data set in Nickell et al. (2005), some observations are still missing. The time series, $\Delta UDNET$ are extended by splicing in data from Nickell et al. (2005). The countries are: Australia before 1965, Austria before 1969, New Zealand before 1972, Portugal in 1975 and 1977, Spain before 1982 and Sweden before 1964. The time series are dividing by 100 to achieve comparable results to the database in Nickell et al. (2005).

The intersection terms between union density and coordination are prolonged by last known observation for these countries.

CO: Coordination of wage setting

The index for coordination of wage setting describes the coordination level in the wage setting. The index range from 1 to 3, and the most coordinated countries have index equal to 3.

The main source is Nickell et al. (2005). The time series are prolonged by the growth rate in the index for coordination in Nickell (2006).

EPL: Employment protection

A measure of the overall employment protection is found in Nickell (2006). Strictness of employment protection is increasing in scale in the range 0 to 2. Se appendix in Nickell (2006) for data definitions, sources and more details.

The time series exists in the period 1960 to 2003. The time series are prolonged backwards and forward by last known observation.

2.B Appendix

This appendix gives additional information to the dynamic simulation of the unemployment rate in section 2.3. The dynamic simulation explores to what extent the empirical model of Nickell et al. (2005) is able to forecast the subsequent post-sample evolution of unemployment, given that we now in general know the correct values of the explanatory variables. The model is evaluated by use of static and dynamic simulation of unemployment for twelve additional years of data. The model does very well in a static simulation. However, as the static simulation is conditional on the lagged unemployment rate, which plays a large role in the model, the static simulation may not give the right impression about the model's out of sample explanatory power. For this reason, the results from the static simulation and some additional dynamic simulations that don't change the main result of the dynamic analysis in section 2.3 are in this appendix, and not in the main text.

In general, the sample period is extended by using the time series for the institutional variables available up to 2003, except for taxes that are available up to 2007, and the time series for the macro variables that are available up to 2007. See appendix 2.A for details regarding the extended and revised data set. The time dummies are set to zero in the simulations. The time trend is either extended or set equal to the average value of the trend and the estimated coefficient from the original data set is used in the simulations. The simulations are also robust to different time dummies extensions: Prolong the time dummies by the last estimated observation or prolong the times dummies by the estimated average of the time dummies (the results are not reported in this paper).

2.B.1 Static simulation of unemployment by using the model in Nickell et al. (2005) on the extended time period 1995 to 2007

The static simulation method simulates unemployment by the empirical model as specified in Nickell et al. (2005), one year ahead out of sample. Within sample period has T periods. This means that we use all available information up to last period, $T+t-1$, to simulate the unemployment rate this period, $T+t$,Clements and Hendry (1998, Ch. 2.7).

Formally, the static simulated unemployment rate for country i can be written as $\tilde{U}^s_{i,T+t} = E(U_{i,T+t}|U_{i,T+t-1}, \mathbf{X}_{i,T+t}, \widehat{\beta})$, where $U_{i,T+t-1}$ is the actual unemployment rate for country i in period $T+t-1$, $\mathbf{X}_{i,T+t}$ is a vector that contains all explanatory variables for country i in period $T+t$ [7] and $\hat{\beta}$ is a vector with all the within sample period estimated

[7] The vector contains mostly actual values of the explanatory variables, but it also contains some predictions for some of the institutions in the period 2001-7, see appendix 2.A for details regarding time periods for the different variables.

coefficients as given in Nickell et al. (2005)˙footnoteThe country specific dummies are taken from the replication in section 2.4.1 The error term is set equal to zero in the simulations.

Figure B1 presents the actual unemployment rate and the static simulated unemployment rate. The model simulates well for Austria, Belgium, Finland, France, Germany, Ireland, Netherlands, Sweden and Switzerland. The simulated unemployment rate is somewhat lower than actual unemployment for 10 countries (Australia, Canada, Denmark, Italy, Japan, New Zealand, Norway, Portugal, United Kingdom and United States) while it is higher for only one country Spain.

As in the dynamic simulation the difference between the simulated and the actual value of the unemployment can be caused by the development in the institutional variables, the specified shock variables or the error term.

Figure B1 also displays the results of a static simulation where the institutional variables are kept constant. Generally, the simulation shows a fairly close fit of the model.

Including institutional variables in the simulation of unemployment generally improves the fit. However, for some countries notable New Zealand, Spain, United Kingdom and United States including institutions actually leads to a worse fit.

The difference between the static simulation and the actual unemployment rate could in principle be due to the development in the specified shock variables. But a simulation where these variables are set to zero after 1995 illustrates that the evolution in static simulated unemployment rate with and without the shocks, essentially have the same development see figure B2.

2.B.2 Dynamic simulation of unemployment by using the model in Nickell et al. (2005) on the extended time period 1995 to 2007

The dynamic simulation in section 2.3 showed a substantial over- and underprediction of the unemployment rate in the post sample period. However, this is not only caused by continuing the time trend. The unemployment rate follows a similar pattern if the trend is prolonged by the last value of the trend in 1995 or by the average value of the trend. The latter gives the largest change in simulations and is shown in figure B3. The simulations shows that the predictive power of the empirical model improves for most countries in the sample, since the simulated value of unemployment by the model is closer to actual unemployment in the post sample period. The prediction of the model improved also in Sweden, but the model now overpredicts the change in unemployment. The predictive power worsens for New Zealand, Germany and United Kingdom.

If the empirical model specified by Nickell et al. (2005) explains the development in unemployment, the difference between the simulated and the actual value of the unemployment can be caused by the development in the institutional variables, the specified

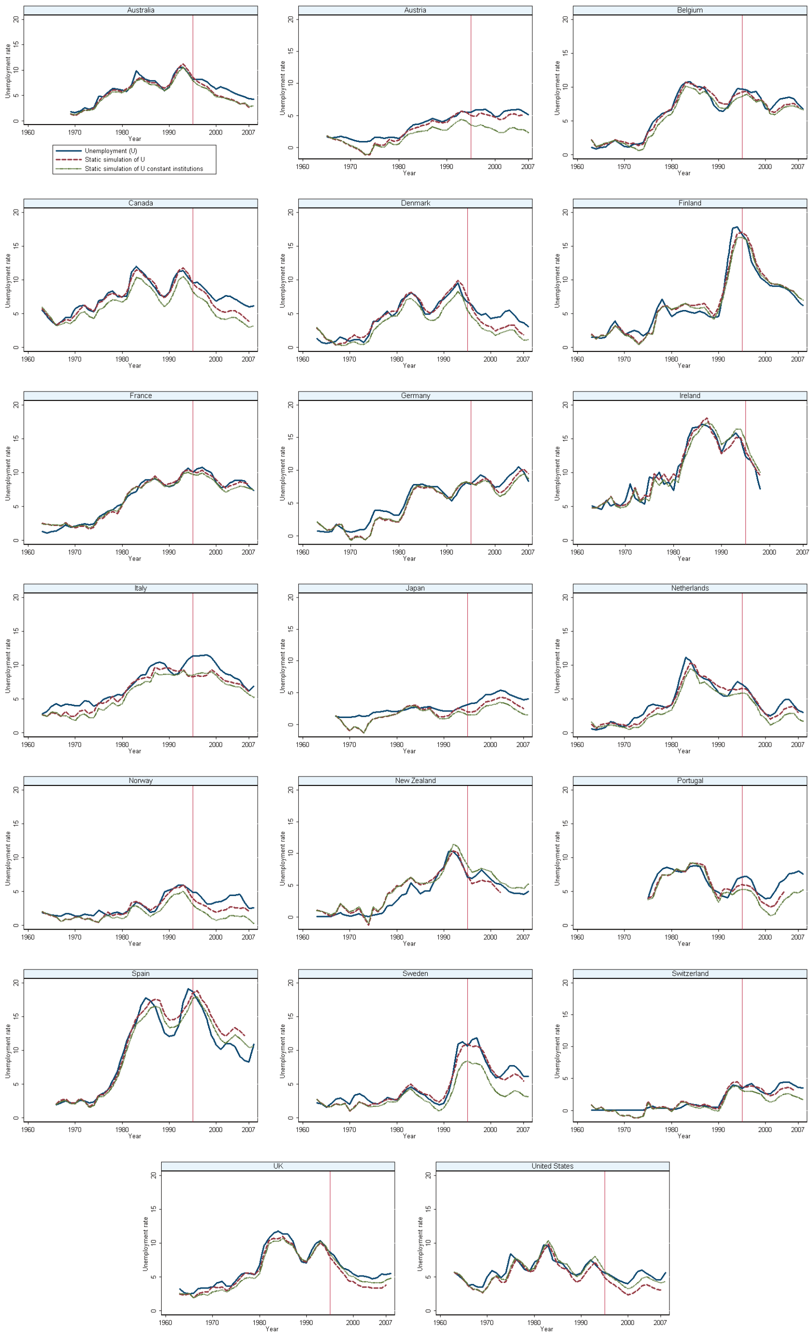

Figure B1: Static simulation and actual unemployment rate. Estimated coefficients on the original data set are used in simulation, in appendix 2.C, model A in table C1. Percent

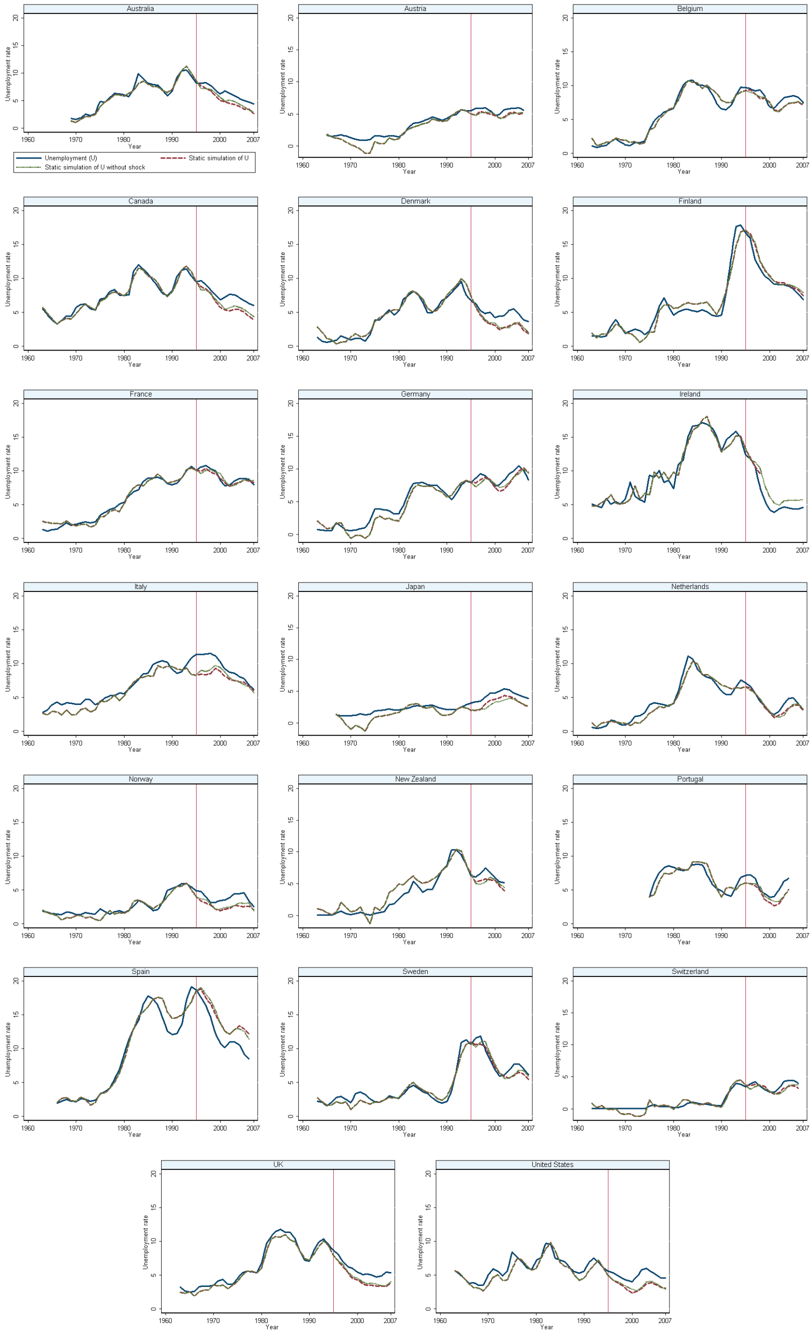

Figure B2: Static simulated and actual unemployment rate. Estimated coefficients on the original data set are used in simulation, in appendix 2.C, model A in table C1. Shocks are zero after 1995. Percent

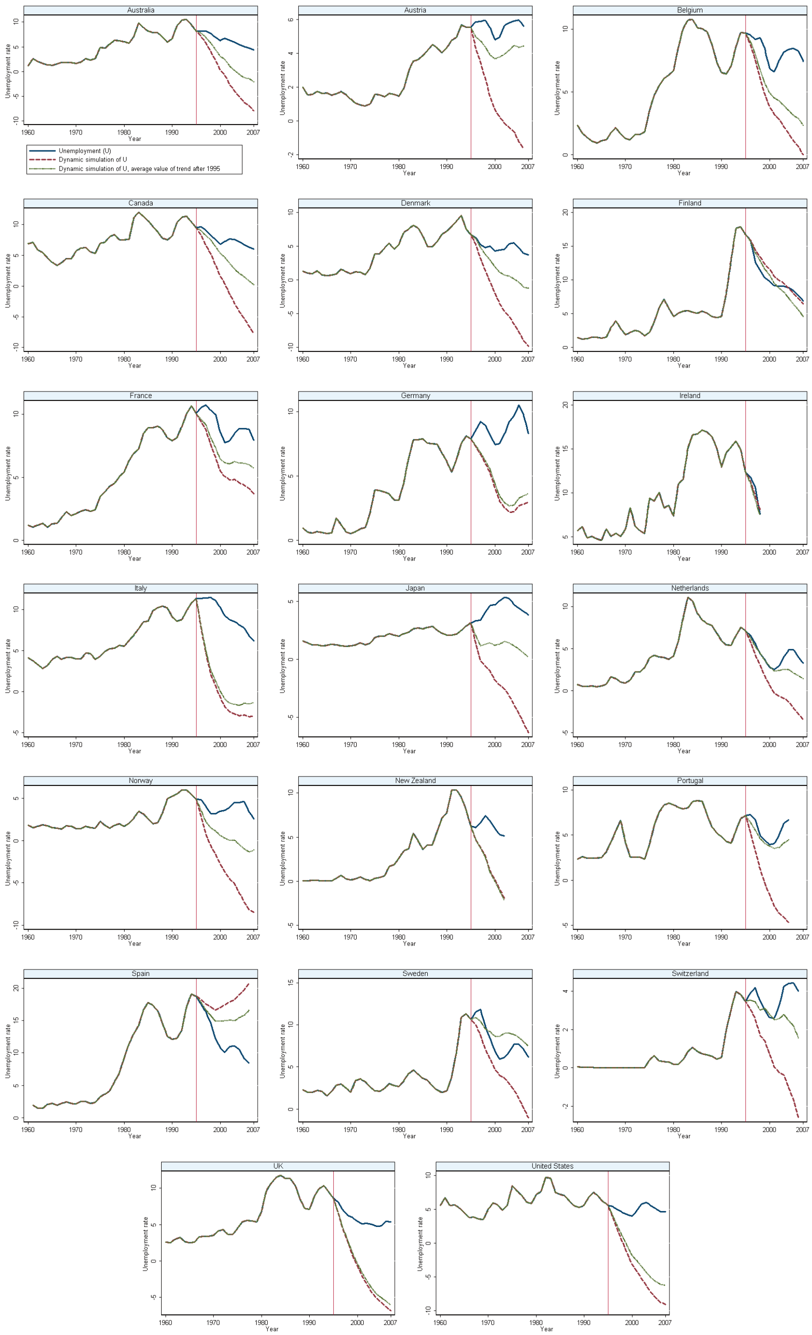

Figure B3: Dynamic simulation and actual unemployment rate. Estimated coefficients on the original data set are used in simulation, in appendix 2.C, model A in table C1. Trend is prolonged with the average value of the trend after 1995. Percent shock variables or the error term. A simulation where the specified shock variables are set to zero after 1995 gives nearly no disparity to the simulation in figure 2, see figure B4.

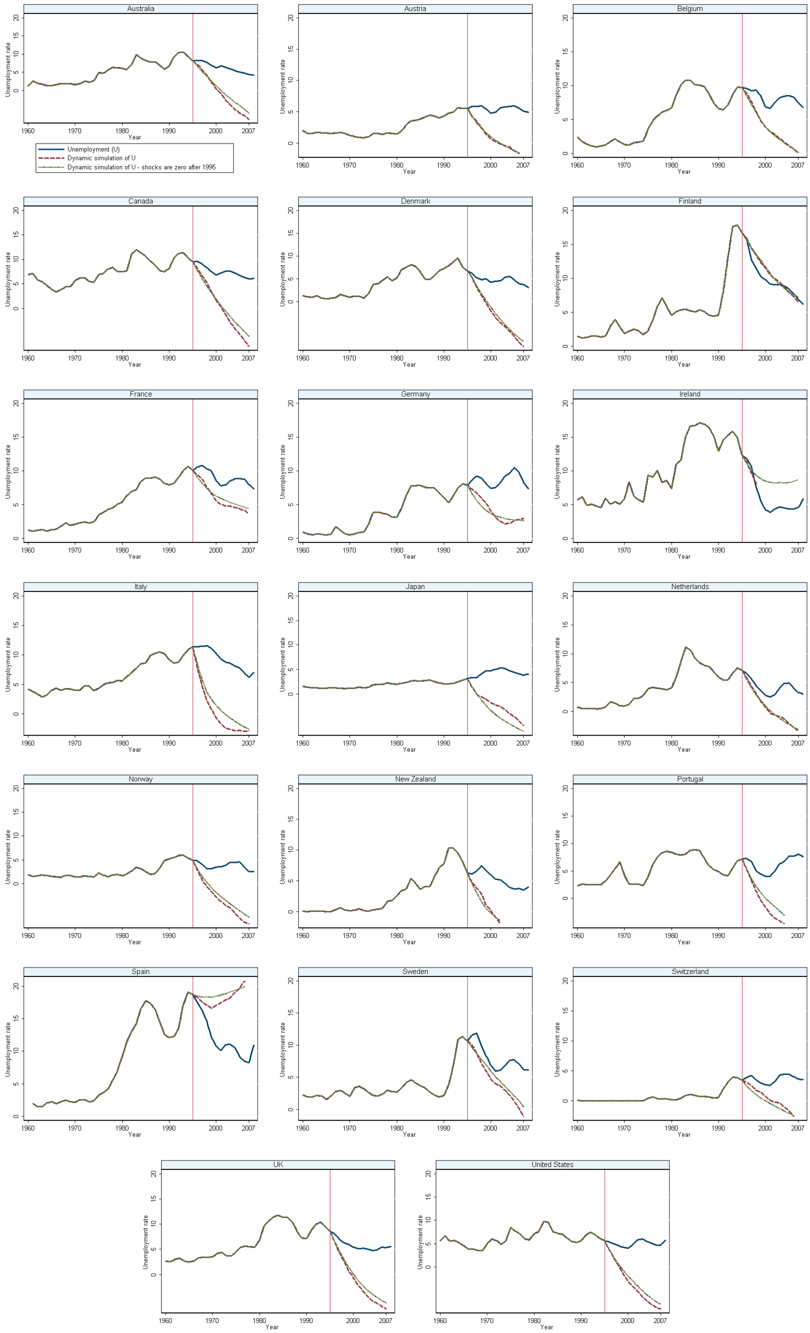

Figure B4: Dynamic simulation and actual unemployment rate. Estimated coefficients on the original data set are used in simulation, in appendix 2.C, model A in table C1. Shocks are set to zero after 1995. Percent

2.C Appendix

This appendix gives some additional information to section 2.4 and 2.5. First, the empirical model in table 5 in Nickell et al. (2005) (NNO hereafter) and of their evaluation by using the original data set[8] is replicated. Second, a static simulation of the model in NNO on the revised and extended data set. Third, a dynamic simulation of the full model, where also the error term specification is taken into account in the simulations on the original data set.

2.C.1 Replication of the NNO model

A replication of empirical model in table 5 in Nickell et al. (2005) and of their results ensures that any differences in results that are found in this paper are due to changes in data revisions or sample length. The estimation procedures and results in Nickell et al. (2005) are described in section 2.2.

I find that the estimated coefficients exactly replicates the coefficient values in Nickell et al. (2005), cf. model A in table C1 and the original results in table 5 in NNO. In addition, a detailed visual inspection of the dynamic simulation on the original data set with and without time-varying institutions are also the same as NNO. The replicated simulations are presented in figure C1.

Table C1: Model A and B is a replication of model 5 in Nickell et al. (2005) on the original data set. In model A, the time dummies are pooled in the time period 1960 to 1966, while the time dummies are pooled in the period 1960 to 1969 in model B

	Model A				Model B			
	Coef.	t-value	min95	max95	Coef.	t-value	min95	max95
Unemployment previous period	0.86	48.49	0.83	0.90	0.86	48.06	0.83	0.90
Employment protection	0.15	0.91	-0.17	0.46	0.16	0.96	-0.16	0.47
Benefit replacement ratio (BRR)	2.21	5.44	1.41	3.00	2.28	5.61	1.48	3.07
Benefit duration (BD)	0.47	2.49	0.10	0.85	0.49	2.66	0.13	0.86
Interaction - BRR and BD	3.75	3.97	1.90	5.60	3.82	4.06	1.98	5.67
Interaction - CO and UDNET	-6.98	-6.12	-9.22	-4.75	-7.04	-6.15	-9.29	-4.80
Interaction - CO and TAX	-3.46	-3.29	-5.52	-1.39	-3.45	-3.29	-5.51	-1.40
First difference in Union density	6.99	3.17	2.67	11.30	7.02	3.19	2.71	11.34
Coordination	-1.01	-3.54	-1.56	-0.45	-0.98	-3.43	-1.54	-0.42
Tax level	1.51	1.72	-0.21	3.24	1.48	1.68	-0.25	3.21
Labour demand shock	-23.58	-10.36	-28.04	-19.12	-23.13	-10.26	-27.55	-18.71
Total factor productivity shock	-17.87	-14.14	-20.35	-15.40	-17.81	-14.33	-20.24	-15.37
Money supply shock	0.23	0.93	-0.25	0.71	0.27	1.14	-0.20	0.75
Real interest rate	1.81	1.56	-0.46	4.08	1.75	1.52	-0.51	4.00
Import price shock	5.82	3.26	2.32	9.33	5.69	3.20	2.20	9.18
Total numb. of obs.	600				600			
Time periods Numb. groups	33	20			33	20		
Log likelihood	-600				-587			
χ^2 of all exogenous variables	46497				45880			

[8] The data is received from Luca Nunziata.

Figure C1: Unweighted Dynamic simulation with the coefficients estimated on the original data set in Nickell et al. (2005) in the period 1960 to 1995. With constant and time-varying variation in institutions. Percent

2.C.2 Static simulation on revised and extended data set

The dynamic simulation explores to what extent the empirical model of Nickell et al. (2005) is able to explain the evolution of unemployment within the sample period when the time series are revised and time period extended.

In general, the sample period is extended by using the time series for the institutional variable available ut to 2003, except for taxes that are available up to 2007. See appendix 2.A for details regarding the revised and extended data set.

In light of the considerable changes in the coefficient estimates in section 2.4 when the time series are revised and extended, one might expect that the model would result in a bad fit of unemployment, at least for the period 1960 to 1995. However, it seems that the new coefficients in a static simulation simulate unemployment well over that period, cf. figure C2. One possible explanation is that the model is quite flexible because of time dummies and country specific trends, in the sense that the model can capture actual unemployment also with the new values of the estimated coefficients of the institutional and shock variables on the revised data set.

However, as the static simulation is conditional on the lagged unemployment rate, which plays a large role in the model, the static simulation may not give the right impression about the fit of the model. For this reason, the results from the static simulation is explored here and the dynamic simulation in section 2.4.

2.C.3 Dynamic simulation of the rewritten model

One possible objection to the empirical specification in Nickell et al. (2005) is that all the heterogeneity is captured by unobserved country and period specific intercepts (a two way error component model), a country specific time trend and a country specific autoregressive error term. In equation (5) in section 2.5, the model is rewritten to explicitly take the effect of the heterogenous lagged error term into the solution of the model for unemployment.

A dynamic simulation of equation (5) with the estimated coefficient values in model A in table C1 above is shown in section 2.5 figure 6 and in figure C3. Figure C3 illustrates a similar pattern as the corresponding graphs in figure C1 which were based on the original model in Nickell et al. (2005). This means that the effect of the heterogeneity in the error term on the simulations is small for a majority of countries, and the two figures show a good correspondence between the two simulations. See further discussion for the countries with unstable solutions related to figure 6 in section 2.5.

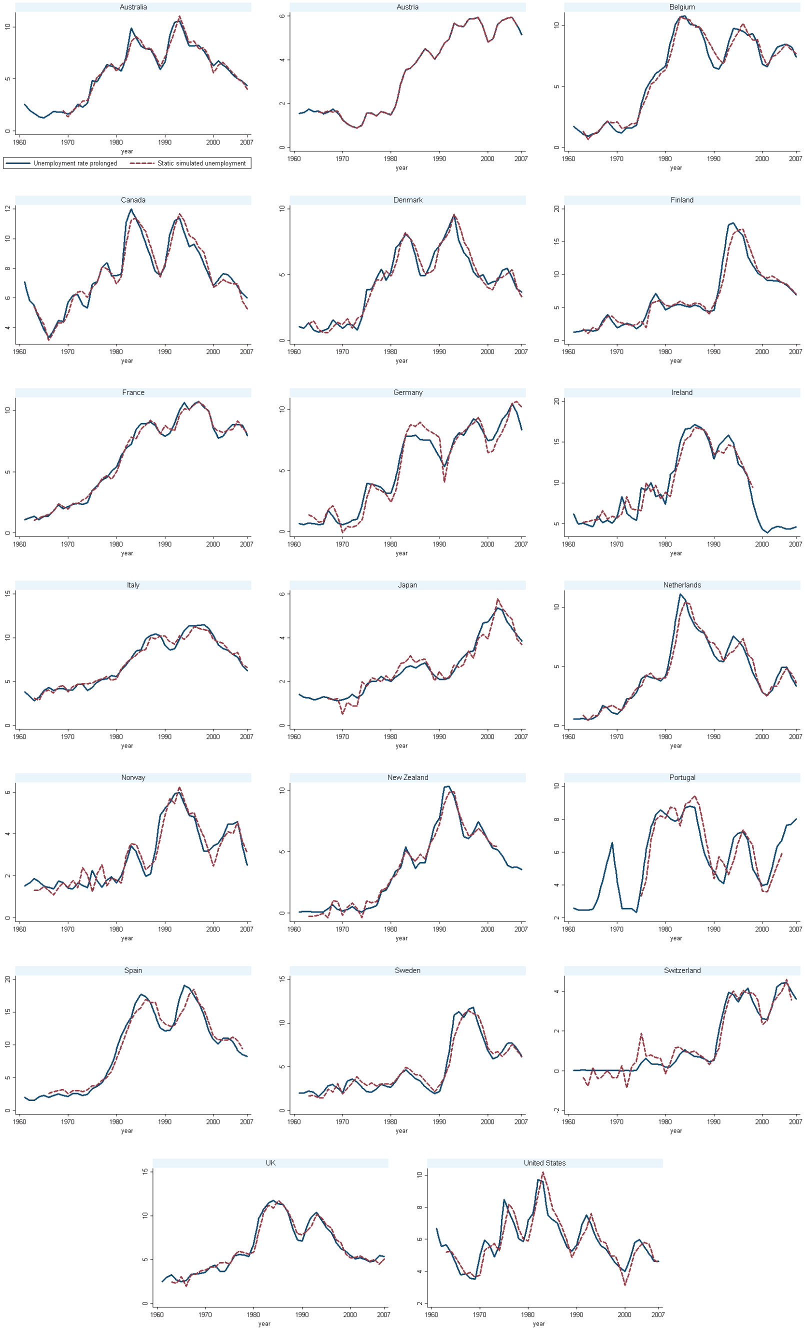

Figure C2: Static simulation of unemployment in the period 1960 to 2007. Estimated coefficients used in simulation are found in table 3. Percent

Figure C3: The transformed equation (5). Model A in table C1. Percent

Chapter 3

The role of institutions in unemployment dynamics and equilibrium.

Ragnar Nymoen and Victoria Sparrman

Abstract We estimate the quantitative importance of labour market institutions for equilibrium unemployment in OECD. The econometric model is based on the solution of a dynamic macroeconomic model which includes structural equations for wage and price setting. Compared to existing studies we use a sample with more variation in unemployment and in institutions, and a higher order dynamics in the final equation for unemployment. Finally, we incorporate objectively and automatically selected indicators for structural breaks. We find that institutional variables have statistical significance, but that these variables account for relatively little of the overall change in the OECD average unemployment rate. The shocks to the economy have been more important for the evolution in the actual average unemployment rate.

We would like to thank Erik Biørn, Neil Ericsson, Øyvind Eitrheim and Steinar Holden for comments and discussions. The numerical results in this paper were obtained by use of *OxMetrics 6/PcGive 13* and *Stata 9*. This paper is part of the project *Demand, unemployment and inflation* financed by the The Research Council of Norway.
University of Oslo, Department of Economics.

3.1 Introduction

The concept of equilibrium unemployment in the OECD area has been subject to both analytical and empirically research. One influential analytical approach, which also underlies our research, combines a model of monopolistic price setting among firms with collective bargaining over the nominal wage level, see Layard et al. (2005). Intuitively, when the system is not in a stationary situation, nominal wage and price adjustments constitute a wage-price spiral that leads to increasing or falling inflation. According to Layard et al. (2005), equilibrium of real wages requires that unemployment becomes equal to the Non-Accelerating Inflation Rate of Unemployment (NAIRU). However, the equilibrium unemployment rate is not interpretable as a constant given from nature. Instead, it depends on different institutional labour market aspects such as wage bargaining coordination, the generosity of the unemployment insurance system and the degree of employment protection. If these institutional variables change, the conflicting real wage aims will change, and the NAIRU will shift.

The panel data literature presents results that support the hypothesis of equilibrium unemployment being affected by the level of labour market institutions for most of the OCED countries. Empirical models for the unemployment rate typically include variables representing the labour market institutions as implied from the above theory, but also interaction between these variables, macroeconomic shocks and interaction between institutional variables and shocks; see Nickell et al. (2005), Bassanini and Duval (2006), Belot and van Ours (2004), Belot and van Ours (2001) and Blanchard and Wolfers (2000). Specifically, Nickell et al. (2005) find a strong role for institutional variables that have an effect on the wage and price setting relationship in explaining unemployment rates in OECD over the period 1960 to 1995. Belot and van Ours (2004) find that specific interactions between labour institutions are the driving forces over the period 1960 to 1999, and Blanchard and Wolfers (2000) find that interactions between institutions and shocks are the main driving forces over the period 1960 to 1995.

In the literature reviewed above, the theoretical framework is static, while the empirical specification of the unemployment equation is either static or dynamic. Dynamics is reasonable given that there are adjustment lags in the manifold of economic, administrative and political decisions that jointly determine the rate of unemployment. The existing studies rely on heuristics to motivate the dynamic specification of the econometric panel data model. Heuristics gives the empirical researcher considerable freedom to choose a specification that fits the data well. One strategy has been to use simple first-order dynamics in the regression and compensate by allowing flexible dynamics in the equations residuals. On the other hand, basing the specification on heuristics alone also means that there is a gap between the underlying theory of equilibrium unemployment, which is static, and the dynamic specification used to estimate the equilibrium rate unemployment.

In this paper we attempt to bridge the gap between the formal, but static, theoretical framework which are common for the studies mentioned, and the dynamic specification of the estimated model based on the framework of Kolsrud and Nymoen (2010). The result of this exercise is a dynamic specification which has some notable differences from the existing ones. First, the estimated dynamic unemployment equation is interpretable as the final equation from a system of equations (the "structural model"). As an implication, the order of unemployment dynamics (first-order, second-order or higher) of the unemployment equation is a consequence of both the number of equations and the order of dynamics in those equations. The third order dynamics is three, and is an extension of earlier papers that use a first-order dynamics (or a second-order dynamics if the residuals are also of first order). A second gain from the formal derivation of the final dynamic equation is that the underlying theory has implication for the signs and the magnitude of the coefficients of the lags of unemployment, which can be confirmed or refuted by estimation. Third, we also get some guidelines for how the labour market variables, which reflect institutional changes, should be brought into the equation. By following the theory and hypothesizing that changes in wage and price settings are primary movers of the equilibrium rate of unemployment, the institutional variables associated with these changes also should enter with potentially long lags. Our specification implements institutions at lags one and two, and shows how the effect of institutions changes with different lag structures. Fourth, the formal derivation of the dynamic unemployment equation also makes it clear that there is no logical or *a priori* reason why the equilibrium unemployment rate cannot be a function of other factors than the labour market institutions. Shocks also from outside the labour market institutions will logically enter the final equation. A change in monetary regime is an example of a break that may have an influence on the equilibrium rate of unemployment, see Iversen (1999) and Holden (2005). In practice, even temporary shocks can be important to include in the model as controls to avoid a bias in the intercept of the equation, which is important for the inferred equilibrium unemployment rate also in panel data models. By allowing for shocks, our formal model is consistent with, for instance, the empirical findings in Blanchard and Wolfers (2000).

We use an econometric algorithm called *Autometrics* as a practical and objective method for determining the shocks that force location-shifts in the unemployment rates, see Doornik (2009). We refer to such shocks as *structural breaks*. In this way, we avoid a specificity problem, namely that one shock which is captured by an included variable may hide the effect of an omitted variable, see Blanchard (2006). Moreover, the structural break variables estimated by our method are interpretable by consulting the economic history of the countries in the data set: Most of the breaks are interpretable as extraneous shocks (the oil price hikes in the 1970s), policy instigated changes (increased interest rates in the US in 1981, and Tatcherism in the UK) or financial crises (the Nordic banking crisis in the early 1990s, see Reinhardt and Rogoff (2009).

In principle all breaks in the rates of unemployment could be due to institutional changes. However, the breaks are uncorrelated with the institutional variables. Another interpretation of our method of detecting shocks is that it is an objective way of controlling for all "special events" which might otherwise have distorted the evidence. Earlier studies have also addressed this problem, but in different ways. Time trends, year dummies and exogenous explanatory variables have been used. Certain countries and years also have been excluded from the sample. Bassanini and Duval (2006) is an example of the latter approach, where Finland, Sweden and Germany are modelled separately due to the shock to these countries caused by the collapse of the Soviet Union.

On the data side, and compared to existing empirical OECD panel data literature, we extend the data set to the period 1960 to 2007. This is important because the time series now contain long periods with both increasing and falling unemployment. The earlier studies in particular over the period 1960 to 1995 may have been dominated by the general increase in the unemployment rate over the same period. As a result, there is more information in the time series dimension than before, which should increase the robustness of the results for the importance of institutions for equilibrium unemployment specifically.

The empirical results is consistent with our theoretical final equation for unemployment. Moreover, the econometric evidence gives support for a role of institutions in the determining of the equilibrium rate, but the quantitative importance is considerably reduced as compared to previous findings in Nickell et al. (2005). This may in part be due to the extended sample, but it is also due to the new econometric specification which has both structural breaks and higher order dynamics. Specifically, if we decrease the most significant variable, the benefit replacement ratio, from its initial average level of 0.5 in 2007 by 20 percent, the OECD average unemployment rate will decrease by 0.8 percentage points. The absence of any large and negative shocks to the economy has been more important for the reduction in the actual average unemployment rate observed before the financial crisis and the following job crisis.

The paper is organized as follows. First, the dynamic model for equilibrium unemployment is derived from the theoretical dynamic model of the wage and price spiral in section 3.2. The data for the evolution of labour market institutions, and the evidence for large shocks (structural breaks), are presented in section 3.3. Econometric issues that are pertinent to the estimation of dynamic models on a macro panel data set are discussed in section 3.4. The results from the estimated dynamic unemployment equations, which include both institutional variables and breaks, are presented and interpreted in section 3.5. We summarize in section 3.6 where we also discuss some extensions and give suggestions for further work.

3.2 A dynamic equilibrium model for the rate of unemployment

To derive a stable final equation for the rate of unemployment, we built on an open economy model as specified in Kolsrud and Nymoen (2010). They focus on the dynamics of the supply side, and close the model by specifying a minimal version of the demand side of the economy which is unimportant for the stability of the wage price spiral, but have important implications for the dynamics of the final equation for unemployment. We therefore alternatively also motivated this specification by starting out from the following identity; the change in the number of unemployed workers from previous period to this period is equal to the flow from employment into unemployment subtracted the new hirings.

Within this framework we specify how changes in institutional variables and more temporary changes in the economic environment can be built into the model. Due to the complexity of the model, we repeat the modelling framework in Kolsrud and Nymoen (2010), but makes explicit comments to the model where and how these variables enter the specification. The dynamic (final) equation for the rate of unemployment follows logically from this model. Under mild assumptions, the equilibrium rate of unemployment can be obtained as the steady-state solution of the estimated equation.

3.2.1 The wage-price spiral

The supply side is modelled such that firms and workers have conflicting interests about the wage share of valued added, created by the joint utilization of capital and labour, within the individual firm as well as in the total economy. However, firms only set the nominal product price, and workers, through wage negotiations, only influence nominal wages. None of the parties have unilateral control over their target real wage variable. This means that when the real wage targeted by the firms is different from the real wage implied by the wage formation, there will be a *wage-price spiral* . The static wage and price curves are interpretable as long-run relationships rather than as equations for actual wage and price setting that hold in each time period.

The starting point of the formal derivation of the wage-price spiral is assumptions for the two exogenous variables import price, pi_t, and productivity a_t. They are both measured in logarithmic scale and are specified as independent random walks with drift:

$$pi_t = g_{pi} + pi_{t-1} + \varepsilon_{pit} \tag{1}$$

$$a_t = g_a + a_{t-1} + \varepsilon_{at} \tag{2}$$

ε_{at}, and ε_{pit} are assumed to be innovations with zero conditional mean. Since domestic wage and price setting are conditional on pi_t and a_t, equations (1) and (2) imply that q_t, the (log of the) price level on domestic products, and w_t, the (log of) wage compensation per hour will be non-stationary, integrated of order one, $I(1)$, in a common notation.

Let p_t denote the logarithm of the consumer price index, and p_t is defined by:

$$p_t = \phi q_t + (1 - \phi) pi_t \tag{3}$$

where the parameter ϕ measures the share of imports in total consumption.

Next define two theoretical (latent) real wage variables: The optimal real wages from the point of view of the firms, rw_t^f, and the bargained real wage, rw_t^b. They are given by the following two equations:

$$\begin{aligned} rw_t^f \equiv w_t^{ef} - q_t^f &= -m_q \qquad\qquad\qquad\;\; + \;\; a_t + \vartheta\, u_t && (4)\\ rw_t^b \equiv w_t^b - q_t^{eb} &= \quad m_w + \omega \left(p_t^{eb} - q_t^{eb}\right) + \iota\, a_t - \varpi\, u_t. && (5)\end{aligned}$$

Equations (4) and (5) are open-economy versions of the relationship for price- and wage-setting in the model originally due to Layard, Nickell and Jackman; see e.g., Layard et al. (2005, p 13).[1]

In the price-setting equation (4), q_t^f denotes the price level set by the firm on basis of expected nominal marginal labour costs $w_t^{ef} - a_t$. w_t^{ef} denotes the *expected* hourly wage cost. We follow custom and assume that pricing is conditional on the actual level of productivity. The last variable in equation (4) is u_t, which is the rate of unemployment (or its log) that is used as a proxy for capacity utilization in this model. The case of $\vartheta = 0$ is called normal cost pricing.

In the wage setting equation (5), w_t^b denotes the nominal wage outcome and q_t^{eb} and p_t^{eb} are the price expectations that affect that bargaining outcome. A main implication of the bargaining model is that elasticity ι with respect to productivity is close to unity, as in Nymoen and Rødseth (1998). The standard assumption about the sign of the coefficient for unemployment ϖ is that it is non-negative, hence $-\varpi < 0$. The coefficient ω is called the wedge-coefficient since $\left(p_t^{eb} - q_t^{eb}\right)$ is the wedge between expected consumer and producer real wages (when we abstract from tax rates). The wedge coefficient is assumed to be non-negative, $\omega \geq 0$.

There is one important difference as compared to the original model of Layard, Nickell and Jackman. They interpret equations (4) and (5) as a model of actual wages and prices, and solve for equilibrium unemployment by invoking the assumption that in equilibrium, there are no expectation errors. The interpretation made here is that neither rw_t^f nor rw_t^b can be set equal to the actual real wage in period t. The reason is that although it can

[1]See also Sørensen and Whitta-Jacobsen (2010, Ch 12 and 17), Blanchard (2009, Ch 6).

be assumed that actual wage setting is influenced by unions and some form of collective bargaining in all countries, *actual* real wages need not be well explained by the simple theory in equation (5).[2]

Instead, the theory of the wage-price spiral is based on the weaker, but not trivial, assumption that rw_t^f and rw_t^b are *co-integrated* with the actual real wage. Similarly, on the price side, it is reasonable that equation (4) captures the secular trend in the actual price q_t, but not the period-to-period changes in the price level.

The co-integration assumption implies real wages to temporary deviate from the equilibrium solution. Temporary deviations are theoretically reasonable if the economy are exposed to shocks and firms not immediate are able (or willing) to adjust the labour force to the new economic environment. Consider the following situation; first wage setters agree upon an real wage which are optimal dependent of the unemployment level. Then the economy is hit by a negative shock which raises the actual wage above the optimal real wage given this particular economic environment. The Layard, Nickell and Jackman model implies an immediately increase in unemployment or inflation. In this model, unemployment rates need not to respond immediate, since the real wages only are assumed to be correlated with the optimal level. This is one assumptions that we could have tested in this paper, but we only consider factors that affect the wage and the price markup, and invoke such changes in the the final equation of unemployment.

We now complete the specification of the dynamic model of wage and price setting: Equations (1)-(4) imply that the two target variables rw_t^b and rw_t^f are $I(1)$ variables. The rate of unemployment, u_t, is assumed to be stationary $I(0)$ after controlling for changes in institutions and other temporary changes in the economic environment. The testing of the empirical relevance of the stationarity assumption is addressed in section 3.3.1 and 3.3.3.

The underlying dynamics of rw_t^b and rw_t^f have important consequences for how the wage-and price dynamics are modelled. First, if the wage and price expectations errors $w_t^{ef} - w_t$, $q_t^{eb} - q_t$ and $p_t^{eb} - p_t$ are $I(0)$, the expectation variables in equations (4) and (5) can be replaced by w_t, q_t and p_t without changing the order of integration. Second, since the non-stationarity of rw_t^b and rw_t^f is an implication of our theoretical model, the theory also implies that rw_t^b and rw_t^f are cointegrated with the actual real wage, rw_t. Third, cointegration implies equilibrium correction, as shown by the Granger-Engle (1987) representation theorem shows. The equilibrium correction model for the wage-price spiral can be written as:

[2] See e.g. the survey evidence from Sweden in Agell and Bennmarker (2007) and the multi-country evidence on wage premiums in Blanchflower and Freeman (1992).

$$\Delta q_t = c_q + \psi_{qw}\,\Delta w_t + \psi_{qpi}\,\Delta pi_t - \varsigma\, u_{t-1} + \theta_q\, ecm^f_{t-1} + \varepsilon_{qt}, \tag{6}$$

$$\Delta w_t = c_w + \psi_{wq}\,\Delta q_t + \psi_{wp}\,\Delta p_t - \varphi\, u_{t-1} - \theta_w\, ecm^b_{t-1} + \varepsilon_{wt}, \tag{7}$$

where ε_{qt}, and ε_{wt} are innovations and all parameters are assumed to be non-negative. The error correction terms, ecm^f_{t-1} and ecm^b_{t-1}, are consistent with equations (4) and (5), with $w^{ef}_t = w_t$, $q^{eb}_t = q_t$ and $p^{eb}_t = p_t$ imposed. They are defined by

$$ecm^f_t = w_t - q_t - \;\; a_t \qquad\qquad - \vartheta\, u_t + m_q \tag{8}$$

$$ecm^b_t = w_t - q_t - \iota\, a_t - \omega\,(p_t - q_t) + \varpi\, u_t - m_w, \tag{9}$$

see Bårdsen et al. (2005) and Bårdsen and Nymoen (2008) for similar derivations.

The dynamic model of the wage-price spiral is:

$$\begin{aligned}\Delta q_t &= (c_q + \theta_q\, m_q) + \psi_{qw}\,\Delta w_t + \psi_{qpi}\,\Delta pi_t - \mu_q\, u_{t-1}\\ &\quad + \theta_q\,(w_{t-1} - q_{t-1} - a_{t-1}) + \varepsilon_{q,t}, && (10)\\ \Delta w_t &= (c_w + \theta_w\, m_w) + \psi_{wq}\,\Delta q_t + \psi_{wp}\,\Delta p_t - \mu_w\, u_{t-1}\\ &\quad - \theta_w\,(w_{t-1} - q_{t-1} - \iota\, a_{t-1}) + \theta_w\,\omega\,(p_{t-1} - q_{t-1}) + \varepsilon_{w,t}, && (11)\\ \Delta p_t &= \phi\,\Delta q_t + (1-\phi)\,\Delta pi_t, && (12)\end{aligned}$$

where equation (12) is the result of taking the difference on both sides of the definition in equation (3).[3] Note that in equations (10) and (11), notations $\mu_q = \theta_q\,\vartheta + \varsigma$ and $\mu_w = \theta_w\,\varpi + \varphi$ are used for the coefficients on u_{t-1}.

3.2.2 Closing the model

In order to close the model, Kolsrud and Nymoen (2010) have referred to a typical medium-term macroeconomic model and assumed that the *GDP output gap* is positively affected by the log of the real exchange rate, $re_t = pi_t - q_t$, and that the unemployment rate is negatively correlated with the output gap, as predicted by *Okun's law*. The mechanism is that an increase in the real exchange rate leads to improved competitiveness. This increases export, and thereby GDP increases and unemployment falls. Based on these conventional assumptions, unemployment is specified as:

$$u_t = c_u + \alpha\, u_{t-1} - \rho\, re_{t-1} + \epsilon_{u,t}, \tag{13}$$

[3]For coefficients ψ_{wq}, ψ_{qw} and ψ_{wp}, ψ_{qpi}, the non-negative signs are standard in economic models. Negative values of θ_w and θ_q imply an explosive evolution in wages and prices (hyperinflation), which is different from the low to moderately high inflation scenario that we have in mind for this paper.

which is a simple, dynamic equation and links the model of the wage-price spiral to the rate of unemployment. As mentioned, an increase in price competitiveness re is assumed to reduce unemployment, so $\rho \geq 0$. It is easiest to think of the current account as the source of this mechanism. The lagged real exchange rate reflect that it takes time before a real depreciation lower unemployment. Note that this assumption is important for the dynamics of the final equation of unemployment, in particular for how changes in institutions enters the equation. We also assume that $-1 < \alpha < 1$ so that this equation alone does not represent a source of *hysteresis* in unemployment, see Røed (1997).

Our alternative interpretation of equation (13) starts with the identity saying that the change in the number of unemployed from $t-1$ to t is equal to the flow from employment into unemployment (separations) minus the flow of new hirings (the number of vacancies filled). In discrete time, we may write as a "law of motion" or, for the unemployment rate u_t

$$u_t = s + (1-s)u_{t-1} - f_t v_{t-1} \tag{14}$$

where v_t is the vacancy rate and s is the separation rate. f_t denotes the rate at which vacancies are filled. In principle, because of discretion and since we have abstracted from population growth, the equation includes a residual which we omit for simplicity.

A key idea in search theory is that f_t depends on the properties of the matching function, see e.g. Pissarides (2000). In the case of a homogenous matching function, we can define $f_t(u_{t-1}/v_{t-1})$ with derivative $f'_{u/v} \geq 0$ and an elasticity between 0 and 1. Linearization then gives

$$f_t \approx f_0 + f_u u_{t-1} + f_v v_{t-1},\ f_u > 0 \text{ and } f_v < 0 \tag{15}$$

The vacancy rate function is probably highly complex, as it depends on the (present) value of opening a new vacancy. For simplicity, it is here written as:

$$v_t = g_0 + g_1\, re_t, \text{ with } g_1 > 0 \tag{16}$$

so that when the "value of the firm" increases due to improved price competitiveness more vacancies are opened.

Inserting equations (15) and (16) in equation (14) gives a relationship like that of equation (13) but the coefficients are interpreted as $c_u = s - f_0 g_0 - f_v g_0^2$, $\alpha = (1 - s - f_u g_0)$ and $\rho = 2 f_v g_0 g_1 + f_0 g_1$. The error term $\epsilon_{u,t}$ in equation (13) then becomes a composite term which mops all the residuals from the approximations in equations (15) and (16), as well as the cross-product $f_u g_1 u_{t-1}\, re_{t-1}$ and the exchange rate squared $f_v g_1^2 re_{t-1}^2$.

In both interpretations, equation (13) is deliberately simple to allow a closed form solution of the full model (10)-(13).

For our empirical specification, we keep in mind that both monetary policy changes and other important policy and institutional differences and reforms can be reflected in the intercept c_u. In the matching model interpretation, for example, we see that since we have $c_u = s - f_0 g_0 - f_v g_0^2$, autonomous changes in both vacancies (g_0) and the matching function (f_0) will influence c_u. This is one of the reasons why, in the econometric panel data model below, we allow for both country-specific intercept dummies and objectively estimated location shifts in the unemployment rates.

3.2.3 The VAR representation

The simultaneous equation system (10)-(13) can be expressed as a vector autoregressive model, VAR:

$$\mathbf{y}_t = \mathbf{R}\,\mathbf{y}_{t-1} + \mathbf{P}\,\mathbf{x}_t + \epsilon_t. \tag{17}$$

Vector $\mathbf{y} = (re, prw, u)'$ contains the endogenous variables and vector $\mathbf{x} = (\Delta pi, \Delta a, a_{-1}, 1)'$ contains the non-modelled exogenous variables and 1 (for the constant term), and vector ϵ contains the reduced form disturbances. $\mathbf{R}$ is a 3×3 matrix and the 3×5 matrix $\mathbf{P}$ contains the coefficients of the non-modelled variables.

The endogenous variable prw_t in the $\mathbf{y}$ vector is the productivity corrected real wage: $prw_t = w_t - q_t - \iota a_t$. With $\iota = 1$ we see that prw_t is simply the log of the wage share: $ws_t = w_t - q_t - a_t$. Kolsrud and Nymoen (2010) show that $\iota = 1$ is also a necessary condition for asymptotic global stability of the system. We impose $\iota = 1$ in the following.

In order to derive the solution for the rate of unemployment in particular, it is useful to write the VAR more extensively, as

$$\underset{\mathbf{y}_t}{\begin{pmatrix} re_t \\ ws_t \\ u_t \end{pmatrix}} = \underset{\mathbf{R}}{\begin{pmatrix} l & -k & n \\ \lambda & \kappa & -\eta \\ -\rho & 0 & \alpha \end{pmatrix}} \underset{\mathbf{y}_{t-1}}{\begin{pmatrix} re_{t-1} \\ ws_{t-1} \\ u_{t-1} \end{pmatrix}} + \underset{\mathbf{P}}{\begin{pmatrix} e & 0 & -d \\ \xi & -1 & \delta \\ 0 & 0 & c_u \end{pmatrix}} \underset{\mathbf{x}_t}{\begin{pmatrix} \Delta pi_t \\ \Delta a_t \\ 1 \end{pmatrix}} + \underset{\epsilon_t}{\begin{pmatrix} \epsilon_{re,t} \\ \epsilon_{prw,t} \\ \epsilon_{u,t} \end{pmatrix}}. \tag{18}$$

The expressions for the coefficients and error terms in equation (18) are given in Appendix 3.C.

3.2.4 The steady-state solution with equilibrium unemployment

The equilibrium rate of unemployment, u^*, is defined as the long-run mean of u_t in Kolsrud and Nymoen (2010). It is the mathematical expectation of the asymptotic stable solution for u_t, obtained from equation (18). Since the model is linear in parameters, u^* is given from the deterministic version of the system. By assuming that the vector of exogenous variables $\mathbf{x}_t = (\Delta pi_t, \Delta a_t, 1)$ is driving the system consists of constants, $\mathbf{x} = (g_{pi}, g_a, 1)$, and that the characteristic roots of $\mathbf{R}$ are inside the unit-circle, the equilibrium rate is obtained as

$$u^* = -\mathfrak{e}_{ss}\, g_{pi} - \underset{\pm}{\mathfrak{b}_{ss}} g_a + \mathfrak{d}_{ss}. \tag{19}$$

where

$$\mathfrak{c}_{ss} = \rho\,(\theta_q(1-\psi_{wq}-\psi_{wp}) + \theta_w(1-\psi_{qw}-\psi_{qpi}))/(\theta_q\,\theta_w\,\Omega), \tag{20}$$

$$\mathfrak{b}_{ss} = \rho\,(\theta_q - \theta_w\,\psi_{qw})/(\theta_q\,\theta_w\,\Omega), \tag{21}$$

$$\mathfrak{d}_{ss} = c_u\,\omega\,(1-\phi) + \rho\,[m_w + m_q + c_w/\theta_w + c_q/\theta_q])/\Omega. \tag{22}$$

with $\Omega = \omega\,(1-\phi)\,(1-\alpha) + \rho\,(\varpi + \vartheta)$. [4]

We see that the term $\mathfrak{d}_{ss}$ in the expression for u^* depends on the two parameters m_w, from wage setting, and m_q, from price setting. m_w in particular represents the "real wage pressure" that results from wage bargaining for the given levels of unemployment productivity and the wedge. Hence, a change in m_w may be due to changes in wage setting institutions and other labour market related institutions, see Nickell et al. (2005), Bassanini and Duval (2006), Belot and van Ours (2004), Belot and van Ours (2001) and Blanchard and Wolfers (2000). Hence, we can write $m_w = m_w(\mathcal{I})$ where $\mathcal{I}$ is a vector of variables that represent the institutional factors for which we have data, see section 3.3.2 below. For symmetry, we may also set $m_q = m_q(\mathcal{I})$, although the firms' real-wage target may be less institutionally conditioned as m_w, see Bjørnstad and Kalstad (2010).

With $m_w = m_w(\mathcal{I})$ and $m_q = m_q(\mathcal{I})$ in place, it follows that $\mathfrak{d}_{ss}$ in equation (19) can shift as a result of changes in m_{q_t} and/or m_{w_t} that are caused by evolving institutions. Changes that occur elsewhere in the economy can also have an effect on m_w and/or m_q as well, but even if $m_w$and/or m_q are unaffected by a large drop in aggregate demand, or by a change in monetary policy, such changes can have an effect through the parameter c_u, as noted in connection with the interpretation of the unemployment relationship equation (13) above.

In section 3.3.3, we show that for each country in our sample, we are able to identify one or more large shocks, or structural breaks in the unemployment time series, which we denote by $\mathcal{D}$. Hence, by writing $c_u(\mathcal{D})$, it follows from equation (20) that the equilibrium rate of unemployment u^* in equation (19) is conditioned by both institutions, $\mathcal{I}$, and large shocks in vacancies, matching and possibly regime shifts elsewhere in the economy.

3.2.5 Unemployment dynamics

The VAR in equation (18) implies a final equation for u_t where the institutional labour market institutions and temporary changes in the economic environment enters the equation in the following way:

$$u_t = \Upsilon_0 + \Upsilon_1 u_{t-1} + \Upsilon_2 u_{t-2} + \Upsilon_3 u_{t-3} + \Upsilon_4 \mathcal{I}_{t-1} + \Upsilon_5 \mathcal{I}_{t-2} + \Upsilon_6 \mathcal{D}_t + \Upsilon_7 \mathcal{D}_{t-1} + \Upsilon_8 \mathcal{D}_{t-2} + \epsilon_{u,t}, \tag{23}$$

[4] This expression is based on the case with wage and price curves, which implies $\mu_w = \theta_w\,\varpi$ and $\mu_q = \theta_q\,\vartheta$. The solution for the version of the model with wage and price Phillips curves ($\theta_q = \theta_w = 0$) is considered in Kolsrud and Nymoen (2010).

In this equation, $\epsilon_{u,t}$ is a composite term that contains lags Δpi_{t-j} and Δa_{t-j} as well as lags of the error terms in equation (18). For completeness, we give the expression for the "gross distrubance":

$$\begin{aligned}\epsilon_{u,t} = &-(l+\kappa)\epsilon_{u,t-1} + (\lambda k + l\kappa)\epsilon_{u,t-2} - \rho\epsilon_{r,t-1} + \rho\kappa\epsilon_{r,t-2} + k\rho\epsilon_{p,t-2} \\ &- \rho e\Delta pi_{t-1} + \rho(\xi k + e\kappa)\Delta pi_{t-2} - k\rho\Delta a_{t-2}\end{aligned} \tag{24}$$

The coefficients Υ_j $(j = 1, 2, 3)$ are uniquely determined by the theoretical model (10)-(13) and are important for the asymptotic global stability of the rate of unemployment. Specifically, the dynamic solution for the rate of unemployment is globally asymptotically stable and converges to u^* in equation (19), if and only if the roots r of the characteristic equation:

$$r^3 - r^2\Upsilon_1 - r\Upsilon_2 - \Upsilon_3 = 0$$

have a modulus less than one.

The expressions for the three autoregressive coefficients are:

$$\begin{aligned}\Upsilon_1 &= \alpha + \kappa + l \\ \Upsilon_2 &= -[\alpha l(1-\kappa) + \kappa(\alpha + l) + n\rho + \lambda k] \\ \Upsilon_3 &= \alpha\lambda k + \rho(n\kappa - \eta k)\end{aligned} \tag{25}$$

From the assumptions, it follows that the first autoregressive coefficient Υ_1 is positive and that it may well be larger than 1. The second autoregressive parameter is expected to be negative, since all terms inside the brackets are positive by assumption. Theory therefore gives clear predictions about the sign and magnitude of Υ_1 and Υ_2. We note that it is possible that $\Upsilon_1 > -\Upsilon_2$, since the additional terms in Υ_2 are products of factors that are less than one. The third autoregressive coefficient, Υ_3, is likely to be smaller in magnitude than the first two coefficients: $\alpha\lambda k$ is a small number and $\rho(n\kappa - \eta k)$ may be near zero or even negative. Therefore, it seems reasonable that the coefficient of the third lag of unemployment may be difficult to discover empirically with a finite amount of data. On the other hand, if the effect is tiny, the bias from estimation of a model with "too little dynamics" may not be significant in numeric terms either.

The coefficients Υ_j $(j = 4, 5, 6, 7, 8)$ depend on how $\mathcal{I}_t$ affects m_w and m_q, and that $\mathcal{D}_t$ primarily affects c_u. The exact expressions for the terms in equation (23) are given below.

A null hypothesis that institutional evolution has no lasting effects on the unemployment equilibrium can be formulated as:

$$\Upsilon_4 + \Upsilon_5 = 0$$

This hypothesis can be tested empirically. The institutional variables enter equation (23) in the following way:

$$\Upsilon_4 \mathcal{I}_{t-1} = \rho d_{t-1} \tag{26}$$

$$\Upsilon_5 \mathcal{I}_{t-2} = k\rho\delta_{t-2} - \rho\kappa d_{t-2} \tag{27}$$

since parameters d and δ are functions of m_q and m_w which, in turn, depend on institutions as explained above. The exact dating of the institutional variables in equation (23), where institutions enter the equation with one and two lags, can be justified by the following reasoning: Changes in institutions in period $t-1$ affect unemployment and wage- and price setting in the next period, t. Another reason, which is also consistent with the above theory, is that mark-ups are affected first and then unemployment. The distributed lag would then be in terms of $t-2$ and $t-3$. That said, a rejection of the null hypothesis of no long-run effect of institutional changes on unemployment should be robust to the exact distributed lag of the institutional variables.

Finally, the dynamics of the VAR in (18) implies that $\mathcal{D}_t$ follows a distributed lag:

$$\Upsilon_6 \mathcal{D}_t = c_{u_t} \tag{28}$$

$$\Upsilon_7 \mathcal{D}_{t-1} = -\left[l + \kappa\right] c_{u_{t-1}} \tag{29}$$

$$\Upsilon_8 \mathcal{D}_{t-2} = l\kappa c_{u_{t-2}} \tag{30}$$

The two interpretations of equation (13) given above, implies that changes in the matching function or simple demand shocks enter the final equation of unemployment with an immediate effect, and with two lags.

However, if shocks also affects the m_w or m_q, in equations (8) and (9), the dynamic effects on unemployment may be different from the dynamics in ($D1$) to ($D3$).

3.3 Data

In this section, we present the evolution in the variables used in our panel data study; unemployment rate, labour market variables and shocks. The panel consists of 20 OECD countries over the period 1960 to 2007. The countries are listed in table 1.

3.3.1 The unemployment rate

The standardized unemployment rate in Economic Outlook at OECD (2008a) is used as a primary data source for the unemployment rate in the OECD countries, see appendix 3.A for details.

There was a substantial change in the unemployment rates of the OECD countries in the period 1960 to 2007. Figure 1 shows the unemployment rates in all countries, together

Table 1: Average unemployment in the OECD countries. Percent

Country	**1960-64**	**1965-72**	**1973-79**	**1980-87**	**1988-95**	**1996-01**	**2002-07**
Australia	1.75	1.79	4.66	7.70	8.41	7.33	5.31
Austria	1.70	1.42	1.38	3.25	4.89	5.49	5.67
Belgium	1.48	1.48	4.23	9.61	8.05	8.34	8.05
Canada	6.00	4.76	6.98	9.84	9.53	8.11	6.92
Denmark	1.07	1.04	3.56	6.48	7.50	5.00	4.62
Finland	1.41	2.41	4.14	5.17	10.85	11.54	8.34
France	1.18	1.95	3.71	7.67	9.10	9.66	8.48
Germany	0.69	0.86	3.06	6.56	6.94	8.31	9.29
Ireland	5.32	5.82	8.08	14.05	14.68	7.30	4.47
Italy	3.46	4.17	4.87	7.96	9.91	10.81	7.71
Japan	1.34	1.24	1.84	2.52	2.46	4.22	4.62
Netherlands	0.57	1.26	3.57	8.28	6.60	4.29	4.03
New Zealand	0.08	0.29	0.74	3.95	8.14	6.37	4.12
Norway	1.71	1.53	1.74	2.44	5.13	3.69	3.89
Portugal	2.46	3.91	5.63	8.23	5.48	5.23	6.90
Spain	1.78	2.31	4.04	14.51	15.00	13.61	9.76
Sweden	2.11	2.61	2.62	3.59	6.22	9.06	6.92
Switzerland	0.03	0.01	0.29	0.63	2.24	3.30	3.99
UK	2.79	3.40	4.81	10.44	8.77	6.31	5.10
United States	5.72	4.47	6.51	7.75	6.16	4.63	5.27
Total	2.14	2.34	3.82	7.03	7.80	7.13	6.17

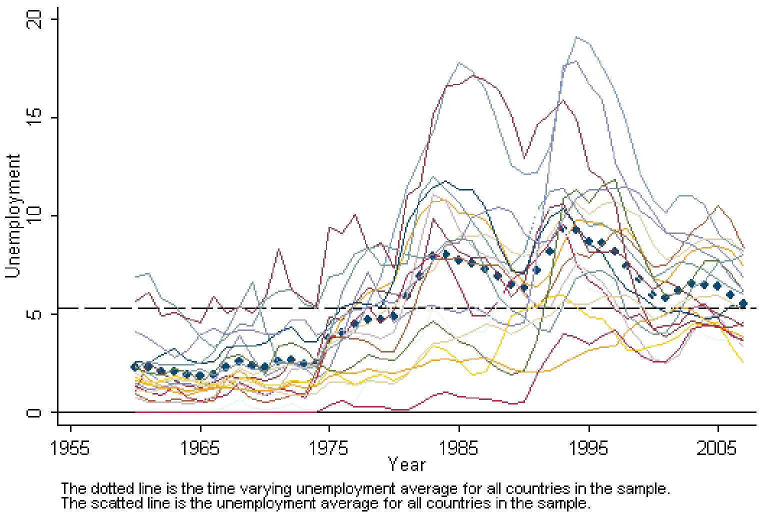

Figure 1: Unemployment rate in the OECD countries. Percent

with the average unemployment rate. The figure illustrates that the rise in unemployment in the early 1970s went together with an increase in the dispersion. The difference between the highest and the lowest unemployment rate in 1995 is larger than in 1960. After 1995, both the average unemployment rate and the variation in unemployment rates across countries have decreased. In our model, institutions and structural breaks capture this development in dispersion.

The evolution of the unemployment rate for each country in the sample in the period 1960 to 2007 is also summarised in table 1. For instance, Norway and Switzerland have a relatively low unemployment rate throughout the period compared to most other countries in the sample. One might question the accuracy of the data showing low unemployment rates in New Zealand and Switzerland in the beginning of the sample period. However, our main results are robust to the exclusion of these countries. Ireland and Spain, on the other hand, are examples of countries with high levels of unemployment in some years and high volatility over time. Germany and Japan have an upward sloping trend, i.e. there is a steady increase in the unemployment rate over time. No country in the data has a tendency of a declining trend in unemployment.

The last row in table 1 contains the average unemployment rate among the OECD countries. Unemployment was very low at the beginning of the period, but increased sharply and peaked at the beginning of the 1980s and 1990s. The average unemployment rate fell slightly towards the end of the period.

The trendlike behavior of the series makes it relevant to test for unit roots. Absence of a unit root is of course essential, since otherwise the empirical relevance of equilibrium unemployment as we have defined it can be questioned, as it has been in the "hysteresis" literature, see Røed (1997).

The Dickey-Fuller tests are the standard tests for unit-root in panel data, see Mátyás and Sevestre (2008). Assume that the unemployment rates are generated by the following first-order autoregressive process:

$$u_{it} = (1 - \alpha_i)\mu_i + \alpha_i u_{it-1} + \epsilon_{it} \tag{31}$$

where the initial values, u_{i0}, are given, and the errors ϵ_{it} are identically, independently distributed across i and t. Equation (31) can be re-written as a Dickey-Fuller regression:

$$\Delta u_{it} = \phi_i \mu_i + \phi_i u_{it-1} + \epsilon_{it} \tag{32}$$

where $\phi_i = \alpha_i - 1$.

The null hypothesis of a unit-root is

$$Ho \quad : \quad \phi_1 = ... = \phi_N = 0$$

Table 2: Unemployment: Dickey-Fuller unit root tests

Alternative:	Homogenous $H1_a$ Statistic (p-value)	Heterogenous $H1_b$ Statistic (p-value)
lag 1	-4.28 (0.00)	-2.29 (0.01)
lag 1, $(U_{it} - \overline{U}_i)^a$	-4.34 (0.00)	-4.04 (0.00)
lag 2	-2.18 (0.02)	-0.45 (0.33)
lag 2, $(U_{it} - \overline{U}_i)^a$	-1.81 (0.03)	-1.99 (0.02)
lag 3	-2.82 (0.00)	-0.91 (0.18)
lag 3, $(U_{it} - \overline{U}_i)^a$	-1.67 (0.05)	-1.80 (0.04)

a) Unemployment (U) subtracted country specific mean ($\overline{U}_i$)

and the two alternative hypotheses are

$$\begin{aligned} H1_a &: \quad \phi_1 = ... = \phi_N \equiv \phi, \text{ and } \phi < 0 \\ H1_b &: \quad \phi_1 < 0, \ ..., \ \phi_{N_0} < 0, \ N_0 < N. \end{aligned}$$

Under $H1_a$ (homogenous alternative), it is assumed that the autoregressive parameter is identical for all cross section units. Under $H1_b$ (heterogenous alternative), it is assumed that N_0 of the N panel units are stationary with individual-specific autoregressive coefficients. A method for testing the null hypothesis against the first alternative is developed by Levin et al. (2002) and a method for the second heterogenous hypothesis is described in Im et al. (2003).

Table 2 shows the test results of the two methods. The tests are performed for three types of augmentations of the basic Dickey-Fuller regression in equation (32), where one, two or three lags in the change in the unemployment rate are included. The table also includes the test statistics which is similar to the above procedures, but first the country-specific mean of unemployment is subtracted from the unemployment rates prior to deriving the test statistic. The procedure mitigates the effect of cross-sectional dependence. “Homogenous alternative” in table 2 shows that the hypothesis of non-stationary is rejected at the conventional five percent significance level for all variations of the test. The “Heterogenous alternative” rejects non-stationarity when the data generating process includes one lag of the change in unemployment, and when the unemployment time series are adjusted for a country-specific mean prior to deriving the test statistic.

Note that the test could depend on whether we control for breaks in the unemployment rate time series as claimed in Camarero et al. (2006). However, since the tests reject the null hypothesis of a unit root without the structural breaks or institutions, the test will also significantly reject the null hypothesis if breaks and institutions are included.

Table 3: Average tax rate in the OECD countries

Country	**1960-64**	**1965-72**	**1973-79**	**1980-87**	**1988-95**	**1996-01**	**2002-07**
Australia	0.23	0.26	0.31	0.35	0.34	0.36	0.37
Austria	0.43	0.48	0.50	0.52	0.52	0.54	0.54
Belgium	0.47	0.51	0.51	0.53	0.59	0.60	0.61
Canada	0.30	0.37	0.38	0.40	0.46	0.47	0.44
Denmark	0.43	0.53	0.53	0.58	0.59	0.63	0.63
Finland	0.41	0.44	0.50	0.54	0.64	0.69	0.68
France	0.66	0.60	0.56	0.60	0.63	0.66	0.64
Germany	0.45	0.45	0.47	0.47	0.47	0.48	0.49
Ireland	0.31	0.35	0.37	0.40	0.42	0.44	0.49
Italy	0.39	0.37	0.36	0.43	0.53	0.64	0.64
Japan	0.24	0.25	0.27	0.31	0.34	0.34	0.35
Netherlands	0.39	0.43	0.48	0.50	0.47	0.45	0.48
New Zealand	0.32	0.32	0.29	0.32	0.40	0.37	0.38
Norway	0.51	0.55	0.57	0.58	0.56	0.59	0.58
Portugal	0.2	0.25	0.26	0.33	0.40	0.44	0.48
Spain	0.20	0.25	0.31	0.40	0.45	0.46	0.51
Sweden	0.38	0.50	0.61	0.69	0.74	0.76	0.75
Switzerland	0.16	0.17	0.21	0.21	0.21	0.24	0.24
United Kingdom	0.34	0.39	0.37	0.40	0.39	0.39	0.40
United States	0.28	0.29	0.30	0.30	0.31	0.32	0.30
Total	0.36	0.39	0.41	0.44	0.47	0.49	0.50

3.3.2 Institutional factors

The main hypothesis to be tested is whether the equilibrium rate of unemployment has been affected by changes in labour market institutions over the sample period. Institutional changes are measured by indices for employment protection (EPL), benefit replacement ratio (BRR), benefit duration (BD), union density ($UDNET$), tax level (TW) and the degree of coordination of wage setting (CO). These indicators are assumed to be correlated with m_{wt} above. We do not have any indicators for m_{qt} or c_{ut}. Appendix 3.A contains a detailed description of all variables and their sources.

The tax rates are calculated from actual tax payments. The total tax wedge is equal to the sum of the employment tax rate (t1), the direct tax rate (t2) and the indirect tax rate (t3). As seen in table 3, there has been a steady increase in tax rates in most OECD countries over the period 1960 to 2007. The largest increases are found in Sweden, Spain and Portugal, and the highest tax rates, larger than a 50 percent tax wedge, are found in the Nordic countries and in Austria, Belgium, France, Italy, Spain and Germany. In Canada and Germany, there was a small decline in the tax rate towards the end of the sample period.

The time series for employment protection measure the strictness of the employment protection for the employee. The overall indicator for employment protection is measured on a scale from 0 to 5. Strictness is increasing in scale. Some other measures only consider the employment protection for regular or temporary contracts. Table 4 shows an average

Table 4: Average employment protection in the OECD countries

Country	**1960-64**	**1965-72**	**1973-79**	**1980-87**	**1988-95**	**1996-01**	**2002-07**
Australia	0.54	0.54	0.54	0.76	0.90	1.20	1.20
Austria	2.20	2.20	2.20	2.20	2.20	2.20	1.95
Belgium	3.32	3.32	3.43	3.30	3.20	2.37	2.20
Canada	0.80	0.80	0.80	0.80	0.80	0.80	0.80
Denmark	2.01	2.01	2.01	2.16	2.19	1.40	1.40
Finland	1.46	1.78	2.30	2.30	2.22	2.08	2.00
France	3.89	4.11	4.11	3.35	2.93	3.00	3.00
Germany	3.52	3.58	3.65	3.43	3.17	2.60	2.25
Ireland	0.63	0.63	0.75	0.89	0.90	0.90	1.07
Italy	3.85	3.91	3.94	3.77	3.60	2.83	1.94
Japan	1.97	1.97	2.07	2.16	2.12	1.90	1.80
Netherlands	2.35	2.35	2.67	2.74	2.70	2.40	2.10
New Zealand	0.90	0.90	0.90	0.90	0.90	1.10	1.50
Norway	3.95	3.95	3.33	2.92	2.88	2.67	2.60
Portugal	4.19	4.19	4.19	4.19	3.95	3.70	3.57
Spain	3.80	3.80	3.80	3.80	3.63	2.97	3.10
Sweden	3.89	3.82	3.55	3.49	3.12	2.25	2.20
Switzerland	0.73	0.76	1.02	1.10	1.10	1.10	1.10
United Kingdom	0.51	0.52	0.57	0.60	0.60	0.63	0.70
United States	0.00	0.00	0.00	0.10	0.20	0.20	0.20
Total	2.23	2.26	2.29	2.25	2.16	1.91	1.83

Table 5: Average benefit replacement ratio in the OECD countries

Country	**1960-64**	**1965-72**	**1973-79**	**1980-87**	**1988-95**	**1996-01**	**2002-07**
Australia	0.18	0.15	0.22	0.23	0.26	0.24	0.19
Austria	0.16	0.16	0.28	0.34	0.34	0.40	0.40
Belgium	0.39	0.37	0.55	0.51	0.48	0.46	0.44
Canada	0.40	0.40	0.60	0.57	0.58	0.52	0.50
Denmark	0.36	0.53	0.79	0.78	0.74	0.66	0.65
Finland	0.20	0.25	0.37	0.43	0.60	0.52	0.35
France	0.47	0.51	0.46	0.59	0.58	0.60	0.61
Germany	0.43	0.42	0.39	0.39	0.38	0.37	0.38
Ireland	0.21	0.24	0.39	0.51	0.41	0.35	0.36
Italy	0.10	0.07	0.04	0.02	0.09	0.45	0.66
Japan	0.36	0.37	0.33	0.28	0.30	0.36	0.40
Netherlands	0.32	0.64	0.65	0.67	0.70	0.70	0.71
New Zealand	0.39	0.30	0.27	0.30	0.29	0.29	0.32
Norway	0.12	0.12	0.24	0.53	0.62	0.63	0.65
Portugal	0.00	0.00	0.14	0.39	0.65	0.66	0.70
Spain	0.17	0.50	0.58	0.74	0.68	0.63	0.64
Sweden	0.24	0.31	0.70	0.85	0.87	0.77	0.74
Switzerland	0.17	0.12	0.27	0.50	0.67	0.72	0.74
United Kingdom	0.27	0.35	0.34	0.27	0.22	0.20	0.19
United States	0.22	0.23	0.28	0.31	0.25	0.29	0.29
Total	0.26	0.30	0.40	0.46	0.49	0.49	0.50

Table 6: Average benefit duration in the OECD countries

Country	**1960-64**	**1965-72**	**1973-79**	**1980-87**	**1988-95**	**1996-01**	**2002-07**
Australia	1.02	1.02	1.02	1.02	1.02	1.01	1.00
Austria	0.00	0.00	0.58	0.75	0.75	0.73	0.78
Belgium	1.00	0.99	0.78	0.79	0.77	0.78	0.80
Canada	0.32	0.33	0.19	0.24	0.23	0.35	0.39
Denmark	0.45	0.49	0.60	0.62	0.69	0.97	0.90
Finland	0.00	0.04	0.65	0.60	0.51	0.66	0.80
France	0.30	0.24	0.20	0.35	0.49	0.50	0.62
Germany	0.57	0.57	0.61	0.61	0.61	0.70	0.81
Ireland	0.67	0.78	0.45	0.38	0.55	0.75	0.75
Italy	0.00	0.00	0.00	0.00	0.00	0.25	0.40
Japan	0.00	0.00	0.00	0.00	0.00	0.04	0.17
Netherlands	0.03	0.35	0.49	0.65	0.60	0.60	0.67
New Zealand	1.02	1.02	1.02	1.03	1.04	1.01	1.00
Norway	0.00	0.02	0.43	0.49	0.50	0.56	0.60
Portugal	0.00	0.00	0.00	0.07	0.35	0.51	0.66
Spain	0.00	0.00	0.00	0.19	0.26	0.31	0.35
Sweden	0.00	0.00	0.04	0.05	0.05	0.05	0.04
Switzerland	0.00	0.00	0.00	0.00	0.07	0.29	0.30
United Kingdom	0.89	0.63	0.54	0.69	0.71	0.80	0.84
United States	0.08	0.16	0.19	0.16	0.19	0.21	0.20
Total	0.32	0.33	0.39	0.44	0.47	0.55	0.60

decline in the strictness of employment protection since the beginning of the 1970s.

The benefit replacement ratio is a measure of how much each unemployed worker receives in benefits from the government in the first period when being unemployed. As seen from table 5, there has been a steady increase in the average benefit ratio for the first period in the period 1960 to 2007. There are large differences in unemployment benefits between the OECD countries, the lowest benefits are found in Australia and the United Kingdom to the highest benefits in Sweden and Switzerland in period 2002-07. Some countries have reversed the benefits during the period, see for instance Canada, Denmark and the United Kingdom, while for instance the United States has been on a low level in the whole sample period.

Benefit duration is a measure of the unemployment benefits for recipients who have been unemployed more than one year, relative to benefits during the first year. As seen from table 6, in many countries the payments stops after one year and the index is then equal to zero. If benefits are the same for the first four years of unemployment, the value of the index is equal to one. If benefits increase over time, the index is larger than one. We observe that in most countries benefit duration has increased over the sample period.

We are interested in the effect of coordination on unemployment, exploring whether centralized wage setters induce wage moderation to reduce unemployment. We use the index in OECD (2004) which measures the formal level of coordination and not whether the coordination actually results in wage moderation at all times as, for instance, in the

Table 7: Average coordination in the OECD countries

Country	**1960-64**	**1965-72**	**1973-79**	**1980-87**	**1988-95**	**1996-01**	**2002-07**
Australia	4.00	4.00	4.04	4.28	2.50	2.00	2.00
Austria	5.00	5.00	4.96	4.35	4.00	4.00	4.00
Belgium	4.00	4.00	3.69	3.96	4.07	4.48	4.50
Canada	1.00	1.00	2.26	1.15	1.00	1.00	1.00
Denmark	5.00	5.00	4.83	3.53	3.40	3.97	4.00
Finland	5.00	5.00	4.91	4.45	5.00	5.00	5.00
France	2.00	2.00	2.00	2.00	2.00	2.00	2.00
Germany	4.00	4.00	4.00	4.00	4.00	4.00	4.00
Ireland	4.00	4.00	3.74	1.79	3.63	4.00	4.00
Italy	2.00	2.00	2.13	2.83	2.90	3.97	4.00
Japan	4.00	4.00	4.00	4.00	4.00	4.00	4.00
Netherlands	3.00	3.00	3.76	4.28	4.00	4.00	4.00
New Zealand	4.00	4.00	4.00	4.00	1.75	1.00	1.00
Norway	4.50	4.50	4.41	3.95	4.50	4.50	4.50
Portugal	5.00	5.00	4.20	3.08	3.75	4.00	4.00
Spain	5.00	5.00	4.29	3.81	3.13	3.00	3.00
Sweden	4.00	4.00	3.96	3.35	3.00	3.00	3.00
Switzerland	4.00	4.00	4.00	4.00	4.00	4.00	4.00
United Kingdom	3.00	3.00	3.46	1.23	1.00	1.00	1.00
United States	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Total	3.67	3.67	3.68	3.25	3.13	3.20	3.20

Table 8: Average union density in the OECD countries

Country	**1960-64**	**1965-72**	**1973-79**	**1980-87**	**1988-95**	**1996-01**	**2002-07**
Australia	0.48	0.45	0.49	0.46	0.38	0.27	0.21
Austria	.	0.62	0.59	0.53	0.45	0.38	0.34
Belgium	0.40	0.42	0.52	0.52	0.54	0.53	0.53
Canada	0.28	0.30	0.34	0.34	0.33	0.31	0.30
Denmark	0.57	0.59	0.71	0.79	0.76	0.75	0.71
Finland	0.35	0.47	0.66	0.69	0.77	0.77	0.72
France	0.20	0.21	0.21	0.15	0.10	0.08	0.08
Germany	0.34	0.32	0.35	0.35	0.32	0.26	0.22
Ireland	0.49	0.56	0.62	0.61	0.55	0.41	0.35
Italy	0.25	0.32	0.48	0.45	0.39	0.36	0.34
Japan	0.34	0.35	0.33	0.30	0.25	0.22	0.19
Netherlands	0.41	0.39	0.37	0.30	0.25	0.24	0.21
New Zealand	.	0.57	0.62	0.59	0.40	0.23	0.22
Norway	0.60	0.57	0.54	0.58	0.58	0.55	0.55
Portugal	.	.	0.59	0.43	0.26	0.21	0.19
Spain	.	.	.	0.10	0.15	0.16	0.15
Sweden	0.66	0.68	0.75	0.80	0.82	0.81	0.76
Switzerland	0.24	0.25	0.27	0.26	0.23	0.21	0.20
United Kingdom	0.40	0.42	0.49	0.48	0.38	0.30	0.29
United States	0.29	0.27	0.23	0.19	0.15	0.13	0.12
Total	0.38	0.42	0.48	0.45	0.40	0.36	0.34

index from Kenworthy (2001). The development in the coordination index is shown in table 7.

Union density rates are constructed using the number of union members divided by the number of employed. Trade union density rates are based on surveys, wherever possible. Where such data were not available, trade union membership and density in European Union countries, Norway and Switzerland were calculated using administrative data adjusted for non-active and self-employed members by Prof. Jelle Visser, University of Amsterdam. Table 8 shows that union density has declined since the beginning of the 1970s in most countries.

We also investigate the interaction between the following institutional variables: benefit duration and benefit replacement ratio, coordination in wage setting and union density, and coordination and tax level. These interaction terms are measured as the deviations from country specific means. For instance, the interaction between coordination and tax is equal to $(CO - \overline{CO})(TW - \overline{TW})$, where $\overline{CO}$ and $\overline{TW}$ are the country specific mean of that variable.

3.3.3 Structural breaks

In the theoretical section, we showed that also other factors than the institutional variables can enter the final equation of unemployment.

Previous studies offer two ways of handling this feature of the data: First, some studies include theoretically motivated variables that represent short-run changes in the unemployment rates, cf. Blanchard and Wolfers (2000) and Nickell et al. (2005). Examples are the change in the interest rate and residuals from short-run changes in labour demand. A second and a complimentary approach is to give special treatment of some years, or some countries in the data set, on account of significant historical events. One example of this approach is Bassanini and Duval (2006) who model Finland, Sweden and Germany separately, because of the effects of the collapse of the Soviet Union on these economies.

In this paper, we take a third approach; we systematically treat the rate of unemployment as a variable which is subject to intermittent structural breaks that may lead to location-shifts. The argument is that the effect of changes in institutional variables on unemployment are likely to be gradual, and are modelled by relatively long lags in accordance with theory above. We have also shown that it is complementary to this hypothesis that the intermittent but large changes in the unemployment rate from one year to another can be due to other factors than institutions, like extraneous or domestic demand shocks, changes in households' preferences for work and leisure or changes in pro-or counter-cyclical economic policies. On the other hand, if the shocks are permanent they could in principle also capture omitted institutions.

To identify shocks in an objective way, we have used the procedures in *Autometrics* for finding the breaks, see Doornik (2009). For each country, we specified a second-order autoregression, and then used two methods called "large outlier" and "impulse saturation" to estimate the structural breaks. The method of "large outlier" adds dummies for years with significant outliers. "Impulse saturation" first adds dummies for each year and then uses the algorithms for automatic model selection to produce a final model with a smaller set of significant structural breaks. The properties of this class of automatic model selection procedures using *Autometrics* are discussed in Castle et al. (2010) and Hendry and Mizon (2010).

Research shows that it is advisable to use a lower level of significance for "impulse saturation" than for "large outliers", and the breaks used in the following are based on the significance level 5 percent of a "large outlier" and 2.5 or lower for impulse saturation, see Doornik (2009). The result is a relatively small number of break dummies. With "large outliers", there is typically just a couple of break dummies. With "impulse saturation", there are more breaks. Based on table 9, the average number of breaks per country from "impulse saturation" is 5.45. With the "large outlier" approach, the average number of breaks is only 1.65.

The majority of the shocks in table 9 are negative location shifts (higher u_t), but there are also positive shocks, in particular in the results from "impulse saturation". Some of these represent the effects of the well-know housing and credit market booms (for example the UK in 1988 and Norway in 2007). There are also effects of "bubbles" that burst at a later stage, for example in the UK in 1991.

Table 9: Impulse saturation and large outlier

Country	**Impulse saturation**	**Large outlier**
Australia	1975, 1977, 1978, 1982, 1983, 1990, 1991, 1992	1983, 1991
Austria	1975, 1981, 1982, 1983, 1989, 1993, 1999, 2000, 2002, 2006, 2007	1982
Belgium	1975, 1981, 1993, 2001, 2002	1975, 1981
Canada	1970, 1975, 1982, 1991	1982, 1991
Denmark	1975, 1981, 1986, 1994	1994
Finland	1967, 1969, 1976, 1977, 1979, 1991, 1992, 1993, 1997	1991, 1992
France	1975, 1984, 1995, 2000, 2007	2000
Germany	1975, 1981, 1982, 1992	1968, 1992
Ireland	1971, 1972, 1975, 1981, 1983, 1995, 1998	1975, 1981, 1983
Italy	1974, 1986, 1993, 1998	1986
Japan	1975, 1988, 2004	1998
Netherlands	1981, 1982, 1983, 1993	1981, 1982, 1984
New Zealand	1983, 1988, 1991	1991
Norway	1975, 1988, 1989, 2006	1989, 2006
Portugal	1970, 1975, 1987, 1993, 1998	1970
Spain	1980, 1984, 1990, 1992, 1993, 2002	1993
Sweden	1971, 1991, 1992, 1993, 1996, 1998, 2003	1993, 1996
Switzerland	1991, 1992, 1993, 1996, 1998, 2002, 2003	1991
UK	1964, 1973, 1980, 1981, 1988, 1991	1981, 1991
United States	1975, 1980, 1982	1975, 1982

To check whether the location-shifts capture something else than the effects of variation in our institutional variables, we have constructed two ("$N \times T$") data series by using the

Table 10: Correlations between breaks and changes in institutions

	Saturation break		Break by large outlier	
	Coefficent	Observations	Coefficent	Observations
Impulse saturation	1.00	960.00		
Change in employment protection	-0.07	940.00	-0.07	940.00
Change in benefit replacement ratio	0.01	940.00	-0.00	940.00
Change in benefit duration	-0.06	940.00	-0.07	940.00
Change in union density	0.12	870.00	0.04	870.00
Change in coordination	-0.14	940.00	-0.13	940.00
Change in tax rate	-0.09	935.00	-0.09	935.00
Large outlier detection			1.00	960.00
Obs.	960		960	

estimated coefficients for the break-dummies, and calculated the correlations between the two location-shift series and the institutional variables, see table 10.
We observe that the correlation between changes in institutions and breaks is generally low, indicating that the breaks do not capture changes in institutions.

The estimated breaks are interpretable on basis of recent economic history. For example, the years 1981-82 are years with location shifts in the unemployment rates of nine countries (seven European countries plus Australia and the US). These years followed the stagflation in the 1970s, the two oil-price shocks, widespread closures in traditional manufacturing in many OECD countries, and marked the start of an evolution of a post-industrial society in many of these countries. For the US in particular, there is only a few well defined breaks, in line with the perception of the US labour market as a flexible one. The first US break, in 1975, can be interpreted as the effect of the oil-price shock. In our interpretation, the break in 1982 captures the effect of the FED's increase of the interest rate to 20 percent at the beginning of the 1980s. This policy was motivated by the need to curb inflation and much of the effect of the interest rate on inflation "went through" the labour market and the rate of unemployment. Another concentration of breaks (8) occur in the first years of the 1990s. This time, the Nordic countries, Finland and Sweden in particular, were also subject to very large cyclical fluctuations and involuntary sharp increase in unemployment.

3.4 Econometric issues

Our primary interest is to estimate the final equation for unemployment, equation (23), derived in section 3.2. To ensure sufficient variation in institutional indicators that may influence the wage and price mark-ups and therefore also unemployment, we follow Nickell et al. (2005) and use macro panel data that consists of 20 OECD countries; however, we extend the period from 1960 to 2007. The evolution of the variables is described in detail in section 3.3, and the definitions and sources are given in the 3.A.

Equation (23) is rewritten to account for country variation i:

$$u_{it} = \beta_{0i}+\beta_1 u_{it-1}+\beta_2 u_{it-2}+\beta_3 u_{it-3}+\beta_4 \mathcal{I}_{it-1}+\beta_5 \mathcal{I}_{it-2}+\beta_6 \mathcal{D}_{it}+\beta_7 \mathcal{D}_{it-1}+\beta_8 \mathcal{D}_{it-2}+\epsilon_{it} \quad (33)$$

where $i = 1, 2, \ldots 20$ and $t = 1960, 1961, \ldots 2007$. Theoretically, ϵ_{it} is a combination of disturbances in price and wage setting, firms' hiring, and the labour supply. We do not impose any particular error structure from the outset. Instead, we test the residual properties of a given specification of the regression and take note if e.g. a test statistic for residual autocorrelation is significant.

Formally, in panel data terminology, the model in equation (1) has heterogeneity in one dimension of the panel, country-specific shifts, and is referred to as one-way heterogeneity (Baltagi, 2008). The country-specific shifts are unobserved, but may be correlated with the explanatory variables and are therefore modelled by the inclusion of dummies for each country. In the case of correlation, omission of β_{0i} in the model would lead to a bias in the estimation of the parameters of the other explanatory variables. Moreover, since the variables that constitute $\mathcal{D}_{it}$ consist of dummies for country-specific structural breaks (for example large demand shocks), we can also interpret the equation as a model with two-way heterogeneity, i.e. with time effects.

It seems plausible that there is more heterogeneity in the "real world" than what our model is furnished with. However, in this paper, we are only interested in the average equilibrium unemployment, and we believe that the heterogeneity modelled here is sufficient for this purpose.

Estimation of equation (1) by OLS, called the least square dummy variable approach (LSDV), leads to biased and inconsistent estimates in the cross section dimention even if the error term ϵ_{it} is not serially correlated, see (Baltagi, 2008, Ch. 8).

A popular alternative approach to the LSDV approach relies on transforming model (1) to first differences, thus eliminating the individual effect. The procedure is then to instrument for the variables that are correlated with the error term and perform a two-stage least square estimation. In this case, the endogenous variable Δu_{it-1} is, by construction, correlated with $\Delta\epsilon_{it}$ and should be replaced by an instrument. Note also that if the null hypothesis of no autocorrelation is rejected empirically, the endogenous variable of a higher lag order will also be correlated with the disturbances and needs to be instrumented for, and the first available instrument will be of higher order too.

A valid instrument is correlated with the endogenous variable and uncorrelated with the error term. The first available instrument correlated with Δu_{it-1} and uncorrelated with the error term (assuming no autocorrelation), is unemployment in period $t-2$. The instrument can also work as an instrument later periods. The number of instruments is therefore quadratic in t. Note also that differences can be used as instruments. The first available instrument in differences is Δu_{it-2}, but this variable is highly correlated with the second endogenous variable Δu_{it-2}. To avoid multicollinearity problems, the first available instrument is from order four, Δu_{it-4}. The exogenous variables are alternative instruments to lagged values of the endogenous variable both in levels and in differences, and can be used as instruments in all time periods.

The Sargan (1958) test is commonly used, see Mátyás and Sevestre (2008, Ch. 4) in the evaluation of the absence of any (asymptotic) correlation between instrumental variables and the disturbances. However, it should be keep in mind that a significant Sargan test can be a sign of econometric misspecification, see Davidson and MacKinnon (2004, Ch. 8.6 and 8.7).

"Difference GMM" follows from the first differences approach, but instead of using only one of the available instruments above, the method uses all available instruments in levels in a one- and two-step procedure, see Baltagi (2008). The calculations are quite complicated but the intuition is that the one-step estimator is derived by replacing the endogenous right-hand side variable in the first differentiated version of equation (1) with all available instruments in levels in a two-stage least square estimation. In our case, this means replacing Δu_{it-1} by a lagged level of unemployment from $t-2$ and higher. The two-step estimator uses the residuals from the one-step process and calculates the variance for each group in the sample. The empirical variance is then used to adjust the weights of each group in the sample, so that the country with the largest variance has least weight when deriving the estimator. The latter GMM estimator requires no knowledge concerning the initial conditions or the distributions of ϵ_{it} or the fixed effect μ_{it} (Baltagi, 2008). If the error term is white noise, the one- and two-step estimators will be asymptotically equivalent.

However, the use of "Difference GMM" on equation (1) is not straightforward in a panel with many time periods because the number of instruments becomes very large, this problem is known as instrument proliferation see for instance Bowsher (2002) and Roodman (2009). The problem affects both the one-step and the two-step estimators. Disregarding the loss of possible instruments due to multicollinearity problems, the instruments grows quadratic in t, and in our case, the number of available instruments is 47^2. The literature notes three problems with many available instruments, this is also our experience: First, we find that applying the "Difference GMM" method overfits the endogenous variable, and the GMM result approaches the OLS results. Second, the number of sample moments used to estimate the optimal weighting matrix for the identifying moments between the instruments and the errors, $var[z'\epsilon]$, is equal to 47^4. This means that the optimal weighting matrix is uniquely identified. Third, this second problem also causes a bias in the Sargan (1958) and Hansen test, see Roodman (2009) and Bowsher (2002). Our results are in line with these findings; the Sargan test rejects the existence of correlation, but the Hansen test showed an implausibly high rejection with a p-value equal to one.

In line with the recommendations from Roodman (2009), we have reduced the number of moment conditions by restricting the number of instruments in two ways: First, the number of lags available as GMM instruments is reduced to 8. Second, the number of moments to be estimated is reduced by collapsing the instrument matrix, which reduces the number of variances per country.

Despite using these methods to reduce available instruments, the results from GMM estimation show a much too large difference between the LSDV results and the GMM results, i.e. they cannot be interpreted as only being corrections of the finite i bias in the fixed effect model[5]. For instance, the estimated coefficient of the lagged endogenous variable changed by a factor of three. In addition, there was a large change in the numerical values of the estimates of the exogenous variables; for instance, benefit replacement changed sign. The reduction in available instruments is therefore probably not sufficient to avoid the erratic behavior of the estimators in this complex model. Also in this case did the Hansen test show rejection with an implausible high p-value equal to one. Note also that the theory gives no clear guidelines for the choice of how to reduce the number of lags available as instruments before the "collapse" function is applied to the model. We have tried with a different number of lags, but all choices reveal erratic behavior of the estimates and large variations (not reported for space consideration).

Another aspect is that several of the LSDV estimates of the coefficients in equation (1) are insignificant. The implication of applying the "Difference GMM" method on a model with insignificant variables is not clear. In practice, a straightforward implementation of the "Difference GMM" means that we instrument the lagged unemployment rate with many weak instruments. It is not clear how this method will affect the weighted matrix or how the resulting estimates should be interpreted. The standard deviation of the residuals in the estimated model has increased by 100 percent, and clearly worsens the fit of the model. The hypothesis of no residual autocorrelation is not rejected, but this is not a problem for consistency. The relevant statistic for the consistency property is the test for the absence of second-order autocorrelation, and this hypothesis is not rejected.

Our results might also suffer from a weak instrument problem caused by the transformation to first differences when the unemployment rate is close to a random walk, see Mátyás and Sevestre (2008, Ch. 8). Then, past unemployment levels convey little information about future changes in the same variable, and untransformed lags may be weak instruments for transformed variables. If past changes are better predictors for current levels than past levels are of current changes, new instruments are more relevant. In our panel data set, the weak correlation between the lags and differences in unemployment rates can be seen in table 11.

This brief overview shows that several alternative estimators exist to the LSDV estimator, but each of the alternatives has its own problems. The bias related to the fixed-effect model should therefore be judged against problems associated with the instrumental vari-

[5]All results from the "Difference GMM" method are available upon request

Table 11: Correlations between lagged values of unemployment and differences

	Unemployment rate (U)		1st diff. in unempl. rate (DU)	
	Coefficient	Observations	Coefficient	Observations
U this period	1.00	959.00	0.09	939.00
U previous period	0.97	939.00	-0.15	939.00
U two years ago	0.92	919.00	-0.26	919.00
U three years ago	0.85	899.00	-0.29	899.00
U four years ago	0.80	879.00	-0.27	879.00
U five years ago	0.75	859.00	-0.24	859.00
DU this period	0.09	939.00	1.00	939.00
DU prev. period	0.20	919.00	0.46	919.00
DU two years ago	0.23	899.00	0.11	899.00
DU three years ago	0.21	879.00	-0.12	879.00
DU four years ago	0.17	859.00	-0.16	859.00
DU five years ago	0.14	839.00	-0.14	839.00
Obs.	959		959	

able alternatives. On the other hand, the differences in estimators might imply that more than one estimator should be considered in a macro model like equation (1). We have chosen to report the fixed-effect estimator, and the one- and two-step estimator for the reduced model where only significant variables enter as explanatory variables.

3.5 Empirical results

In this section, we start by estimating versions of equation (1) where we allow for other forcing variables than the institutional factors, namely the "structural breaks" that we motivated in section 3.3. First, we discuss the role of institutions within this model. Then we explore whether the results are sensitive to the dynamic specification or the choice of method of detecting structural breaks.

3.5.1 The role of institutions

The column marked "All countries" in table 12 shows the fixed effect estimation results of equation (1) with the sequence of location-shift dummies obtained from the "large residuals" method. The lower part of the table contains the two χ^2-test relevant for the role of institutions. They both reject their respective joint null-hypotheses of no effect of institutions. The value of the test statistic for the significance of levels effects of the institutional variables is 37.63, and the value of the test statistics for the significance of the interaction terms is 22.62. From the detailed coefficient estimates of the different variables, we see that the level of employment protection and the benefit replacement ratio are both statistically significant at the 5 percent level so that both stricter employment protection and higher replacement ratio lead to higher unemployment. For the two interaction terms, only the short-term effects are significant, the change in the interaction between coordination and union density, and the change in interaction between coordination and taxes significantly reduces unemployment. At the significance level of 10 percent there is also evidence of a long-term effect on unemployment from the interaction between benefit replacement ratio and benefit duration, so that a change in this variable increases unemployment.

Table 12: Estimates from the fixed effect model with large outlier dummies

	All countries		All countries[a]		Heterosc.		Red. data set[b]	
	Coef.	Std	Coef.	Std	Coef.	Std	Coef.	Std
Unemployment prev. period	1.38	0.03	1.31	0.03	1.45	0.03	1.38	0.03
Unemployment two years ago	-0.52	0.05	-0.43	0.05	-0.65	0.05	-0.52	0.05
Unemployment three years ago	0.06	0.03	0.02	0.03	0.13	0.03	0.06	0.03
Employment protection (EPL), 1st diff. prev. period	0.12	0.24	-0.11	0.22	0.10	0.21	0.14	0.25
EPL, two years ago	0.14	0.07	0.04	0.07	0.15	0.06	0.14	0.07
Benefit replacement ratio (BRR), 1st diff. prev. period	-0.89	0.81	-0.85	0.73	-0.64	0.69	-0.88	0.84
BRR, two periods ago	0.63	0.24	0.35	0.21	0.55	0.20	0.63	0.26
Benefit duration (BD), 1st diff. prev. period	-0.51	0.54	-0.36	0.49	-0.14	0.43	-0.49	0.56
BD, two periods ago	0.00	0.17	-0.13	0.16	0.07	0.14	0.03	0.18
Interaction - BRR and BD, 1st diff. prev. period	-2.59	2.03	-1.31	1.81	-1.37	1.63	-2.90	2.16
Interaction - BRR and BD two periods ago	1.17	0.63	1.16	0.56	1.05	0.52	1.27	0.68
Interaction - CO and UDNET, 1st diff. prev. period	-4.16	2.15	-3.62	1.92	-3.06	2.35	-4.30	2.43
Interaction - CO and UDNET two periods ago	-0.77	0.46	-0.86	0.41	-1.05	0.36	-0.98	0.53
Interaction - CO and TW, 1st diff. prev. period	-8.02	2.48	-6.05	2.24	-4.63	1.92	-8.40	2.70
Interaction - CO and TW two periods ago	-0.18	0.81	0.10	0.71	-0.19	0.63	-0.05	0.84
Union density (UDNET), 1st diff. prev. period	0.43	2.02	1.84	1.88	-0.18	1.87	0.56	2.29
UDNET, two periods ago	0.26	0.28	0.11	0.30	0.23	0.27	0.39	0.31
Coordination (CO), 1st diff. prev. period	0.12	0.17	0.20	0.16	0.00	0.16	0.18	0.18
CO, two periods ago	-0.01	0.04	-0.01	0.04	-0.01	0.04	-0.00	0.04
Tax rate (TW), 1st diff. prev. period	-0.29	1.52	1.74	1.40	1.67	1.31	-0.05	1.61
TW, two periods ago	0.48	0.53	0.40	0.55	0.63	0.46	0.42	0.57
Break by Large outlier approach	0.94	0.05	0.79	0.04	0.90	0.05	0.94	0.05
Tot. obs and the number of countries	837	20	837	20	837	20	761	18
Standard deviation of residuals	0.6		0.5		0.6		0.6	
χ^2 of all the exogenous variables.[c]	531.93	(0.00)	417.04	(0.00)	451.42	(0.00)	490.27	(0.00)
χ^2 of institutional variables (level).[c]	37.63	(0.00)	31.41	(0.03)	37.33	(0.00)	33.83	(0.01)
χ^2 of institutional variables (interaction).[c]	22.62	(0.00)	20.28	(0.00)	22.52	(0.00)	21.47	(0.00)
1st order autocorrelation[c]	0.37	(0.71)	0.37	(0.71)	0.37	(0.71)	0.51	(0.61)
2nd order autocorrelation[c]	-1.71	(0.09)	-1.71	(0.09)	-1.71	(0.09)	-1.49	(0.14)

a) With time dummies.

b) Without New Zealand and Switzerland.

c) Numbers in parenthesis are p-values for the relevant null.

Table 12 also contains three additional estimation results; "All countries[a]" which includes time dummies for each year which are common to all the countries in the sample, "Heterosc." which is a GLS estimation which accounts for heterogeneity in the error term and "Red. data set" which excludes New Zealand and Switzerland due to the unrealistic low values for unemployment at the beginning of the period, see section 3.3 for data details. Time dummies are included in "All countries[a]" since macroeconomic shocks that are common to all countries in the sample might bias the estimated coefficents. "Heterosc." is included as a robustness test since the theoretical derivation shows that the disturbance term of the model may contain short term influences from changes in the world price and productivity growth, cf. equation (24) above.

For all models in table 12, the conclusions based on the χ^2-test are the same as in the "All countries" model, which is one way of illustrating the importance of institutions. A closer inspection of table 12 reveals that, as a rule, the sign and the significance of the coefficients in "All countries" are retained in all models. The exception is employment protection, which is insignificant in the "All countries^a" model.

When it comes down to results for other individual variables, both the direct effect of union density and the direct effect of tax rates have changed signs in the "Heterosc." model. However, it is not obvious what the correct short-run coefficient is. Several authors have claimed that the causality between institutions and unemployment is unclear: For instance, a higher coordination level can imply lower wage claims if the coordination level is above a certain level, cf. Calmfors et al. (1988). The wage claims are then a function of the degree of coordination, where medium level of coordination results in the highest wage clams. The low and high coordination levels result in low wage claims. A similar argument also applies to union density. Holden and Raaum (1991) argue that increased union density in some cases may facilitate wage moderation and thus induce lower unemployment.

The long-run solution to the estimated model in table 12, "All countries", is presented in equation (34). The numbers in parenthesis below the coefficients are asymptotic standard deviations, and the long-run t-values can be obtained by dividing the estimated coefficients by these standard deviations. The long-term effects of institutions all have the signs that we expect from theory. Benefit duration, coordination and the interaction between benefit replacement ratio and benefit duration, increase unemployment, while the interaction between coordination and union density decreases unemployment. The long-run t-value is larger than two in absolute value for all these variables. The institutional variables which are significant in the long-run equation correspond well to those variables which have low p-values in table 12.

$$\begin{aligned} u^* = {} & Constant + \underset{(0.9)}{1.6} EPL + \underset{(2.2)}{7.6} BRR + \underset{(1.6)}{0.03} BD \\ & + \underset{(2.8)}{3.1} UDNET - \underset{(0.5)}{0.1} CO + \underset{(6.4)}{5.2} TW + \underset{(5.5)}{14.2} (BRR - \overline{BRR})(BD - \overline{BD}) \\ & - \underset{(3.5)}{9.1} (CO - \overline{CO})(UDNET - \overline{UDNET}) - \underset{(6.9)}{3.0} (CO - \overline{CO})(TW - \overline{TW}) \end{aligned} \tag{34}$$

We illustrate the quantitative effect of institutions on the average OECD unemployment rate by two dynamic simulations of the "All countries" model in table 12. Both simulations start in 1969 and end in 2007. The first simulation is conditional on the actual values of all non-modelled exogenous variables over the solution period. We call this the solution with *time varying institutions*. The second simulation is based on constant values (from 1968) of the institutional variables, and we call this the solution with *constant institutions*. The residual term is set to zero in both simulations. The result of the simulation is shown in figure 2.

Overall, the model with "large outlier breaks" seems to fit the data quite well since the gap between the simulated unemployment rate and actual unemployment is small in figure 2. The unweighted average unemployment rate in OECD is estimated to be 6.0 percent in 2007 which is close to 5.5 percent, i.e. the actual value of the average unemployment rate in OECD in 2007. The figure also shows a small gap between the simulation with time-varying and constant institutions, and illustrate that only a small part of the evolution of unemployment can be attributed to changes in institutions. Note however that on average there have been only small changes in the institutional variables over the OECD countries in table 13.

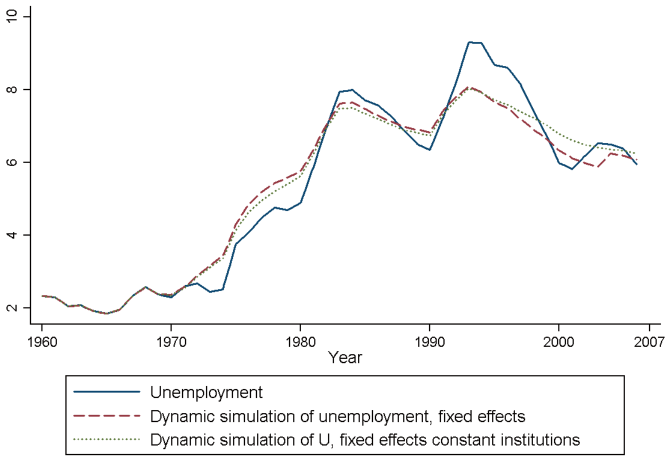

Figure 2: Dynamic simulation of the OECD average unemployment rate. Estimated coefficient values from table 12 "All countries" (break by large outlier). Simulations with and without time varying institutions

Several of the variables in table 12 and also in equation (34) have insignificant estimated coefficients. When we drop all variables that are insignificant at the 10 percent level from table 12 in "All countries" and reestimate the simplified equation, we obtain the more parsimonious model in the column of table 14 labelled "Fixed effects". The results show that the variables which have sizable effects in the general model remain significant and the values of the estimated coefficients are of the same magnitude also in the simplified model. The corresponding long-run equation is:

$$
\begin{aligned}
u^* = {} & Constant + \underset{(0.8)}{2.0}\,EPL + \underset{(1.8)}{7.9}\,BRR \\
& + \underset{(7.5)}{13.5}(BRR - \overline{BRR})(BD - \overline{BD}) - \underset{(4.2)}{9.6}(CO - \overline{CO})(UDNET - \overline{UDNET}) \qquad (35)
\end{aligned}
$$

We can use equation (35) to illustrate the long-run effects of labour market institutions in a different way than the dynamic simulation. One of the most significant variables is the benefit replacement ratio. The average value is equal to 0.5 in 2007. According to equation (35), a reduction in the benefit replacement ratio of 20 percent will decrease average OECD unemployment with 0.8 percentage points. Note that the effect is somewhat stronger if the interaction effect with benefit duration is included. The overall conclusion is that the change in labour market institutions have to be quite large in order to lower the OECD unemployment rate substantially.

Table 13: Actual changes in institutions over the period 1960 to 2007

Institutional variable X:	$\bar{X}^a_{69-60} - \bar{X}^b_{89-85}$	$\bar{X}^c_{07-00} - \bar{X}^b_{89-85}$	$\bar{X}^c_{07-00} - \bar{X}^a_{07-60}$
Employoment protection (EPL)	-0.03	-0.36	-0.39
Benefit replacement ratio (BRR)	0.21	0.01	0.22
Benefit duration (BD)	0.13	0.15	0.28
Interaction BRR BD	-0.04	0.00	-0.04
Interaction CO UDNET	0.00	0.00	0.00
Interaction CO TW	0.01	0.01	0.02
Union density (UDNET)	0.03	-0.08	-0.05
Coordination (CO)	-0.47	0.00	-0.48
Tax rate (TW)	0.10	0.03	0.13

a) $\bar{X}_{69-60}$ is the average level of the institutional variable in the period 1960 to 1969.
b) $\bar{X}_{89-85}$ is the average level of the institutional variable in the period 1985 to 1989.
c) $\bar{X}_{07-00}$ is the average level of the institutional variable in the period 2000 to 2007.

Table 14 also contains the results of the Arellano-Bond one- and two-step estimation method, with lagged levels of unemployment in addition to the exogenous variables as GMM instruments. The main impression is that there are small differences between the results for the two estimation methods, and all variables except for taxes have the same sign (compare the results under "Fixed effects", "Arellano bond, onestep" and "Arellano bond, two step" in table 14). Taken at face value, this shows that the "bias-problem" of the LSDV estimator does not constitute a major issue for the parsimonious model. This is as expected for a sample like ours, where the time series are quite long and there are no roots "on" the unit circle.

Table 14: Estimation result for a simplified model with large outlier dummies, fixed effects and Arellano Bond one-step and two-step.

	Fixed effects			Arellano bond, onestep			Arellano bond, twostep		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment prev. period	1.37	0.02	0.00	1.19	0.06	0.00	1.17	0.04	0.00
Unemployment two years ago	-0.45	0.02	0.00	-0.40	0.03	0.00	-0.43	0.04	0.00
EPL, two years ago	0.11	0.06	0.07	0.40	0.23	0.08	0.46	0.13	0.00
BRR, two periods ago	0.60	0.20	0.00	0.45	0.63	0.48	0.33	0.62	0.60
Interaction - BRR and BD two periods ago	0.98	0.55	0.07	0.35	1.67	0.83	0.45	1.41	0.75
Interaction - CO and UDNET 1st diff. prev. period	-4.59	2.09	0.03	-6.14	2.74	0.03	-5.30	0.83	0.00
Interaction - CO and TW 1st diff. prev. period	-7.58	2.37	0.00	-8.38	2.75	0.00	-4.89	3.14	0.12
Break by Large outlier approach	0.95	0.04	0.00	0.83	0.05	0.00	0.79	0.02	0.00
Tot. obs and the number of countries	913	20		893	20		893	20	
Standard deviation of residuals	0.59			0.52			0.52		
χ^2 of all the exogenous variables.[a]	26.69	(0.00)		17.32	(0.00)		95.17	(0.00)	
χ^2 of institutional variables (level).[a]	10.99	(0.00)		3.23	(0.20)		11.99	(0.00)	
χ^2 of institutional variables (interaction).[a]	17.00	(0.00)		12.64	(0.01)		49.81	(0.00)	
1st order autocorrelation[a]	1.07	(0.29)		-7.57	(0.00)		-2.22	(0.03)	
2nd order autocorrelation[a]	-0.04	(0.97)		-1.97	(0.05)		-0.46	(0.64)	
Sargan test[a]				16.35	(0.00)		16.35	(0.00)	
Hansen test[a]							13.46	(0.04)	

a) Numbers in parenthesis are p-values for the relevant null.

3.5.2 The role of dynamic specification

We now discuss how the empirical conclusions derived in the previous section depend on the dynamic specification and the exact dating of institutions on unemployment.

The autoregressive part of table 12 “All countries” corresponds to a characteristic equation with three roots; one real root is equal to 0.85 and two roots are complex with moduli equal to 0.59. Since all roots are well inside the unit circle, the model has a stable steady-state solution, which is also consistent with the more formal tests of stationarity in section 3.3. The absence of a unit root is of course essential since otherwise, the dynamic stability assumption that underlies the existence of an equilibrium level of unemployment would be empirically unfounded. Also the other models with heterogenous residuals and the results from the reduced data set are consistent with the assumption of a stable long-run mean of the rate of unemployment (conditional on a fixed value of the institutions).

The estimated autoregressive coefficients in table 12 “All countries” correspond well with the *a priori* magnitude derived from plausible assumptions for the model parameters which we discussed in section 3.2: First, the first-order coefficient is large and positive. Second, the coefficient of u_{t-2} is negative and highly significant. And, finally, u_{t-3} is numerically small.

Table 15, model 1 shows the results of an estimation where the autoregressive lags are reduced to one lag, but is otherwise similar to the model in table 12 “All countries”. The value of the estimated autoregressive coefficient is close to one. Reduced dynamics therefore implies more persistence in the evolution of the unemployment rate. All the significant institutional variables in table 12 “All countries” are significant in table 15 in model 1, except for the interaction term between benefit replacement ratio and benefit duration. In addition, benefit duration, the 1st difference in union density, and union density in previous period are significant in table 15 in model 1. However, the first order autocorrelation test in the lower part of the table rejects the hypothesis of no autocorrelation. This is a sign of misspecification and it also damages the formal test based on t-values. In this sense our theoretically motivated dynamic specification is supported by the evidence.

As explained in the theoretical section above, the dynamic specification of the institutional variables is based on the assumption that changes in institutions in period $t-1$ jointly affect unemployment and the wage- and price setting in the next period, t. From another perspective, it might be hypothesized that we put institutional variables at a disadvantage by excluding within year effects. Indeed, the specification without lags has been used in previous literature, e.g. Nickell et al. (2005), Bassanini and Duval (2006) and Blanchard and Wolfers (2000).

The estimation results when we change the dynamic specification of institutions are shown in table 15, model 2 and 3. In model 2, all the institutional variables enter contemporaneously and with one lag, while in model 3 the institutions enter with a level this period. As shown in the lower part of table 15, model 2 and 3, the first χ^2-test for the institutional variables in level rejects the hypothesis of no joint effect from labour market institutions on unemployment. However, the χ^2-test for the interaction terms between the institutional variables are insignificant. The autocorrelation tests for model 2 model reject the hypothesis of 1st. and 2nd. order autocorrelation, while the 1st order autocorrelation is not rejected in model 3.

Equation (36) gives the corresponding equilibrium unemployment equation for model 2 in table 15:

$$\begin{aligned} u^* = {} & Constant + \underset{(1.0)}{1.8} EPL + \underset{(2.1)}{5.5} BRR + \underset{(1.5)}{0.4} BD \\ & + \underset{(3.3)}{5.5} UDNET - \underset{(0.5)}{0.4} CO + \underset{(7.2)}{8.6} TW + \underset{(5.0)}{11.0} (BRR - \overline{BRR})(BD - \overline{BD}) \\ & - \underset{(4.6)}{11.9} (CO - \overline{CO})(UDNET - \overline{UDNET}) - \underset{(8.5)}{6.1} (CO - \overline{CO})(TW - \overline{TW}) \end{aligned} \tag{36}$$

The estimated coefficients in this equation are not too different from the estimated coefficient in equation 34. We conclude that the exact lag specification is of minor importance for capturing the effect of the institutional variables. This is not surprising given that the labour market institutions change gradually.

Table 15 model 4 also shows the results of a static equation. The estimated coefficients in model 4 are completely different from the other models, and the test of the residual autocorrelation at the end of the table confirms an increasing degree of misspecification, since the test statistics indicate both first- and second-order residual autocorrelation.

Table 15: Estimation results for the fixed effect model with large outlier dummies, different dynamic specifications

	Model 1		Model 2		Model 3		Model 4	
	Coef.	Std	Coef.	Std	Coef.	Std	Coef.	Std
Unemployment previous period	0.95	0.01	1.38	0.03	1.39	0.03		
Unemployment two years ago			-0.52	0.05	-0.54	0.05		
Unemployment three years ago			0.06	0.03	0.07	0.03		
Employment protection (EPL), 1st diff. prev. period	-0.08	0.28						
EPL, two years ago	0.28	0.08						
Empl. protection (EPL), 1st difference			-0.36	0.24				
EPL prev. period			0.14	0.07				
EPL this period					0.11	0.07	-0.87	0.24
Benefit replacement ratio (BRR), 1st diff. prev. period	-0.54	0.94						
BRR, two periods ago	0.72	0.27						
Benefit repl. ratio (BRR), 1st difference			-0.66	0.81				
BRR prev. period			0.41	0.24				
BRR this period					0.50	0.23	5.62	0.80
Benefit duration (BD), 1st diff. prev. period	-1.32	0.63						
BD, two periods ago	-0.13	0.19						
Benefit duration (BD), 1st difference			-0.09	0.55				
BD prev. period			0.02	0.17				
BD this period					0.08	0.17	0.42	0.59
Interaction - BRR and BD, 1st diff. prev. period	-2.53	2.37						
Interaction - BRR and BD two periods ago	0.73	0.72						
Interaction BRR and BD, 1st difference			-0.09	2.04				
Interaction BRR and BD prev. period			0.83	0.63				
Interaction BRR and BD this period					0.96	0.62	16.16	2.13
Interaction - CO and UDNET, 1st diff. prev. period	-6.24	2.51						
Interaction - CO and UDNET two periods ago	-1.28	0.53						
Interaction CO and UDNET, 1st difference			0.28	2.15				
Interaction CO and UDNET prev. period			-0.89	0.46				
Interaction CO and UDNET this period					-0.81	0.45	-3.77	1.63
Interaction - CO and TW, 1st diff. prev. period	-9.79	2.87						
Interaction - CO and TW two periods ago	-0.92	0.92						
Interaction CO and TW, 1st difference			-2.87	2.46				
Interaction CO and TW prev. period			-0.31	0.81				
Interaction CO and TW this period					-0.45	0.80	-1.37	2.78
Union density (UDNET), 1st diff. prev. period	4.54	1.87						
UDNET, two periods ago	0.55	0.32						
Union density (UDNET), 1st difference			6.31	2.00				
UDNET prev. period			0.42	0.29				
UDNET this period					0.52	0.28	4.41	0.98
Coordination (CO), 1st diff. prev. period	-0.08	0.20						
CO, two periods ago	0.03	0.05						
Coordination (CO), 1st difference			0.06	0.17				
CO prev. period			-0.03	0.04				
CO this period					-0.02	0.04	-1.34	0.14
Tax rate (TW), 1st diff. prev. period	-2.89	1.74						
TW, two periods ago	-0.27	0.60						
Tax rate (TW), 1st difference			-1.56	1.52				
TW prev. period			0.70	0.54				
TW this period					0.28	0.53	16.30	1.70
Break by large outlier	0.99	0.05	0.92	0.05	0.94	0.04	0.65	0.17
Tot. obs and the number of countries	851	20	838	20	844	20	886	20
Standard deviation of residuals	0.7		0.6		0.6		2.2	
χ^2 of all the exogenous variables.[a]	497.21	(0.00)	528.53	(0.00)	495.32	(0.00)	549.98	(0.00)
χ^2 of institutional variables (level).[a]	70.42	(0.00)	29.54	(0.00)	12.75	(0.05)	495.24	(0.00)
χ^2 of institutional variables (interaction).[a]	25.26	(0.00)	6.81	(0.34)	5.68	(0.13)	61.57	(0.00)
1st order autocorrelation[a]	3.52	(0.00)	0.09	(0.93)	3.48	(0.00)	3.70	(0.00)
2nd order autocorrelation[a]	0.17	(0.87)	-1.61	(0.11)	0.70	(0.49)	3.64	(0.00)

a) Numbers in parenthesis are p-values for the relevant null.

3.5.3 The role of identifying structural breaks

As discussed above, we have used two methods for estimation of location-shift variables: "impulse saturation" and "large outliers". As also noted, the "impulse saturation" approach leaves less variation to be explained by changes in institutions as compared to the "large outlier" approach, simply because the "impulse saturation" approach gives more year dummies. Whether this leads to more or less explanatory power of the included variables remains to be seen. As a benchmark model, we also investigate the model where we exclude the break variables.

In table 16, "All countries" shows OLS results for the fixed effects model with breaks estimation determined by the "impulse saturation" method. Compared to table 12 (result with "large outlier"approach), employment protection and the difference in union density are no longer significant, and the effect of tax rates has changed sign even though the effect of taxes is insignificant both in tables 12 and 16.

The autoregressive part of table 16, "All countries" corresponds to a characteristic equation with three real roots equal to 0.91, 0.27 and 0.12, this implies that also this model has a stable steady-state solution.

The long-run solution to the estimated model in table 16 is:

$$
\begin{aligned}
u^* = {} & Constant + \underset{(1.0)}{1.3} EPL + \underset{(2.2)}{8.2} BRR + \underset{(1.7)}{0.4} BD \\
& + \underset{(3.0)}{0.6} UDNET - \underset{(0.5)}{0.9} CO - \underset{(6.5)}{2.1} TW + \underset{(7.6)}{18.0} (BRR - \overline{BRR})(BD - \overline{BD}) \\
& - \underset{(5.8)}{12.8} (CO - \overline{CO})(UDNET - \overline{UDNET}) - \underset{(6.8)}{5.6} (CO - \overline{CO})(TW - \overline{TW}) \qquad (37)
\end{aligned}
$$

We observe that the benefit replacement ratio alone, the interaction with benefit duration, and the interaction between coordination and union density all have significant effects on unemployment. The long-run effects correspond well with the variables that have low p-values in table 16. Compared to the long-run solution in equation (34), the significant estimated coefficients have nearly the same magnitude in both equations, while the insignificant variable taxes, has changed sign.

Figure 3 shows the average dynamic simulated unemployment rate of equation (1) with the estimated coefficient values from "All countries" in tables 12, 16 and B1, where the latter table is found in appendix 3.B. The motivation for bringing in the appendix result is that this model is estimated without any location-shift variable, hence the corresponding simulated solution is denoted "without breaks" in figure 3. The model with "impulse saturation" has a visually better fit than the model with "large outlier", and the estimated unweighted average unemployment rate in OECD with "impulse saturation" is 5.8 percent and closer to the actual average unemployment rate than the model with dummies from the "large outlier" data series.

A simulation with and without timevarying institutions illustrates the quantitative

Table 16: Estimates from the fixed effect model with saturation breaks

	All countries		All countries[a]		Heterosc.		Red. data set[b]	
	Coef.	Std	Coef.	Std	Coef.	Std	Coef.	Std
Unemployment prev. period	1.30	0.02	1.26	0.02	1.35	0.02	1.30	0.02
Unemployment two years ago	-0.39	0.04	-0.34	0.04	-0.49	0.04	-0.39	0.04
Unemployment three years ago	0.03	0.02	-0.00	0.02	0.08	0.02	0.03	0.02
Employment protection (EPL), 1st diff. prev. period	0.10	0.19	0.00	0.17	0.16	0.17	0.10	0.19
EPL, two years ago	0.08	0.05	0.05	0.05	0.10	0.05	0.08	0.05
Benefit replacement ratio (BRR), 1st diff. prev. period	-0.90	0.63	-0.88	0.59	-0.19	0.53	-0.90	0.63
BRR, two periods ago	0.52	0.19	0.34	0.17	0.45	0.16	0.52	0.19
Benefit duration (BD), 1st diff. prev. period	-0.18	0.43	-0.07	0.40	-0.21	0.33	-0.18	0.43
BD, two periods ago	0.02	0.13	-0.10	0.13	0.13	0.10	0.02	0.13
Interaction - BRR and BD 1st diff. prev. period	-2.01	1.59	-1.63	1.46	-0.29	1.29	-2.01	1.59
Interaction - BRR and BD two periods ago	1.16	0.49	1.06	0.45	1.23	0.43	1.16	0.49
Interaction - CO and UDNET 1st diff. prev. period	-6.73	1.69	-6.08	1.56	-4.70	1.87	-6.73	1.69
Interaction - CO and UDNET two periods ago	-0.84	0.36	-0.92	0.33	-1.33	0.28	-0.84	0.36
Interaction - CO and TW 1st diff. prev. period	-4.43	1.95	-4.00	1.81	-2.35	1.53	-4.43	1.95
Interaction - CO and TW two periods ago	-0.19	0.63	-0.09	0.58	0.01	0.49	-0.19	0.63
Union density (UDNET), 1st diff. prev. period	-0.31	1.58	1.19	1.52	0.33	1.45	-0.31	1.58
UDNET, two periods ago	0.03	0.22	0.08	0.24	0.08	0.19	0.03	0.22
Coordination (CO), 1st diff. prev. period	0.07	0.13	0.13	0.12	-0.06	0.13	0.07	0.13
CO, two periods ago	-0.06	0.03	-0.05	0.03	-0.06	0.03	-0.06	0.03
Tax rate (TW), 1st diff. prev. period	0.74	1.19	1.83	1.13	1.84	1.01	0.74	1.19
TW, two periods ago	-0.09	0.42	-0.23	0.45	0.08	0.35	-0.09	0.42
Saturation break	0.93	0.03	0.83	0.03	0.90	0.03	0.93	0.03
Tot. obs and the number of countries	837	20	837	20	837	20	837	20
Standard deviation of residuals	0.5		0.4		0.5		0.5	
χ^2 of all the exogenous variables.[c]	1397.41	(0.00)	1079.91	(0.00)	1401.92	(0.00)	1397.41	(0.00)
χ^2 of institutional variables (level).[c]	43.96	(0.00)	47.17	(0.00)	59.27	(0.00)	43.96	(0.00)
χ^2 of institutional variables (interaction).[c]	33.42	(0.00)	33.71	(0.00)	44.04	(0.00)	33.42	(0.00)
1st order autocorrelation[c]	0.62	(0.53)	0.62	(0.53)	0.62	(0.53)	0.74	(0.46)
2nd order autocorrelation[c]	-0.30	(0.76)	-0.30	(0.76)	-0.30	(0.76)	-0.17	(0.87)

a) With time dummies.
b) Without New Zealand and Switzerland.
c) Numbers in parenthesis are p-values for the relevant null.

effect of institutions on the average OECD unemployment rate. The importance of the institutional changes for the development of unemployment appears to be stronger with the “impulse saturation”, i.e. with the larger number of breaks, since the gap between simulated unemployment with and without time-varying institutions is visually larger in figure 4 than in figure 2. This result is achieved despite of the fact that the first model leaves less of the variation to be explained by institutions. This could illustrate the importance of controlling for other factors influencing unemployment to achieve the true effect of institutions.

The results of the long-run steady-state projections are presented in figure 5. The models “All countries” in tables 12, 16 and B1, where the latter is found in appendix 3.B, extending into the future the end-of-sample values of the institutional variables and assuming no future location-shifting breaks in the rate of unemployment. This gives some insight into the speed of unemployment adjustment. It also gives a picture of the implied equilibrium level of unemployment, u^*, based on the assumptions just mentioned. The steady-state solution of the estimated dynamic model in table 12 is, in practice, determined by simulation, keeping the institutional variables fixed at their 2007 level, and by switching off the “large outlier breaks”. The simulated unemployment rate will then converge to a steady state. The effect of the “large outlier breaks” influences the estimates of the institutional variables and the autoregressive parameter, even though

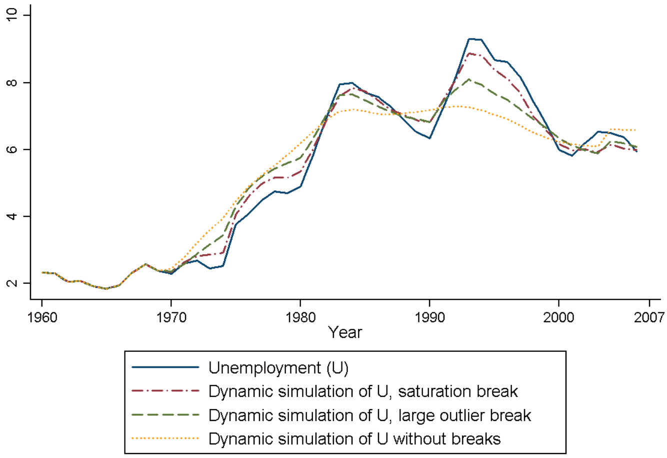

Figure 3: Dynamic simulation of unemployment by using the estimates in tables 16, 12 and B1, "All countries", where the latter is found in appendix 3.B

they only have a temporary effect.

The graphs in figure 5 also show that even controlling for shocks that are impulses rather than step-functions is important for the estimated level of equilibrium unemployment. Intuitively, when the structural breaks explain a larger part of the growth in unemployment, the simulation until 2037 when no structural breaks are imposed, leads to lower unemployment. The figure illustrates that "impulse saturation" gives the lowest estimate for equilibrium unemployment, then comes "large outlier" and the highest level is the model without any dummy included.

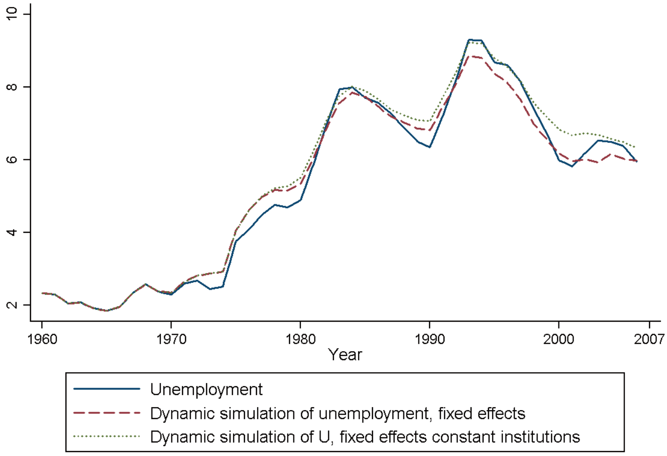

Figure 4: Dynamic simulation coefficient values from the estimation in table 16, "All countries" (saturation break) with and without time varying institutions, unweighed average of the unemployment rate in OECD

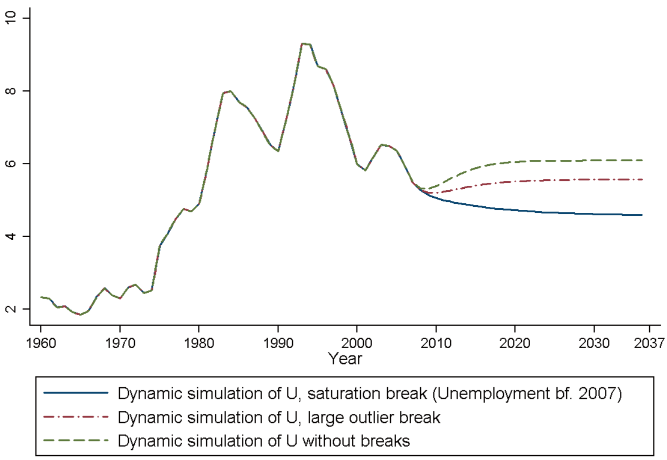

Figure 5: Dynamic simulation of estimates in tables 16, 12 and B1, where the latter is found in appendix 3.B from 2007 to 2037. The unweighed average of the unemployment rate in OECD, within sample

3.6 Conclusions

The equilibrium rate of unemployment is an important parameter in economic models that are used for forecasting and as an aid for policy analysis and decisions. A realistic estimate of the equilibrium rate could, for example, be used together with other targets to assess the economic performance of the economy, and motivate medium-term fiscal and monetary policies. Also when another variable than unemployment is targeted (i.e. an inflation target), one priority of policy may be to avoid large destructive shocks to the economy which threaten to shift the equilibrium rate upwards through increased long-term unemployment and discouraged worker effects, see Layard et al. (2005). Furthermore, it may be an aim for more long-term economic policy to reduce the equilibrium rate of unemployment through institutional and other structural reforms. These are some of the reasons for trying to capture the main determinants of the equilibrium rate of unemployment and model the dynamics of the actual rate of unemployment around that equilibrium.

The results in this paper confirm earlier findings that labour market institutions have a significant effect on the OECD unemployment rate. The estimated long-term effects of the institutions all have the signs as expected from theory (maybe with one exception, a increase in the union density leads to higher unemployment, but the effect is not significant). The parsimonious long-run solution, where all variables with a lower significance level than 10 percent are dropped from the model, shows that unemployment is increasing in the following variables: level of employment protection, the unemployment benefit replacement ratio, and the interaction between benefit replacement ratio and benefit duration. An increase in the interaction between coordination and union density decreases unemployment. These results imply that appropriate institutional reforms have the potential of lowering equilibrium unemployment.

That said, if one looks at the magnitude of the estimated coefficients, labour market institutions have to change quite substantially to achieve a sizeable reduction in the unemployment rate. For instance, the isolated effect of a reduction in benefit replacement ratio of 20 percent, from the OECD average in 2007, will lower the average OECD unemployment rate by 0.8 percentage points. This effect is half the size of the estimated long-run effect of the unemployment benefit replacement ratio that was found in Nickell et al. (2005). The small effect of institutions is consistent with historical evidence, a simulation of our main unemployment equation with and without time varying institutions reveal a small effect of institutions over the sample period, but one should be aware of the fact that on average there have been small changes in the institutional variables over the sample period.

Over the same period, we find that the actual rates of unemployment have reacted to shocks in a way that has dominated the evolution of the unemployment rate. We suggest to treat the rate of unemployment as a variable which is subject to intermittent structural breaks, that may lead to location-shifts. The argument is that the effect of changes in institutional variables on unemployment are likely to be gradual, and are modelled by relatively long lags in accordance with theory. The intermittent but large changes in the unemployment rate from one year to another is treated as caused by other factors than institutions, like extraneous or domestic demand shocks, changes in households' preferences for work and leisure, or changes in pro-or counter-cyclical economic policies. We call these changes structural breaks. We have chosen two statistical methods of detecting such shocks, "large outliers" and "impulse saturation".

The inclusion of structural breaks that capture location shifts in the distributions for the unemployment rates turns out to be important for our estimate of the equilibrium rate. If we do not correct for these structural breaks, the equilibrium rate is simulated to almost 6.2 percent, while the lowest adjusted estimate is 4.3 percent. However, comparing the simulation of the two models with structural breaks shows that the model with more breaks illustrate a larger gap between time varying and constant institutions. This might illustrate the importance of controlling for other factors influencing unemployment in order to achieve the true effect of institutions.

In terms of modelling methodology, this paper has also illustrated the importance of the dynamic specification of the panel data model for the rate of unemployment. We show that a reduced lag structure on the autoregressive coefficients increases the residual autocorrelation, which might be a sign of misspecification. The chosen dynamic specification is derived theoretically and has the status of a final equation of a system consisting of equations for wage and price setting and an equation of unemployment as a function of the real exchange rate. The theoretical derivation gives *a priori* assumptions regarding the magnitude of the autoregressive coefficients. The assumptions are confirmed by the empirical evidence. On the other hand our results also show that the exact lag structure of the institutional variables are of minor importance for capturing the effects of labour market institutions on unemployment.

In this paper, the shocks have been identified using the automatic model specification for each country. An interesting extension and improvement of this methodology is to use the panel dimension also in the identification of the structural breaks. Another extension is to provide interval estimates for equilibrium rate of unemployment.

3.A The Data: Definitions and sources

This appendix contains information about variables that are important for the evolution of the unemployment rate in 20 OECD countries. The countries in the sample are:

Australia	Finland	Japan	Spain
Austria	France	Netherlands	Sweden
Belgium	Germany	Norway	Switzerland
Canada	Ireland	New Zealand	United Kingdom
Denmark	Italy	Portugal	United States

The variables in this data set are divided into two groups; economic variables and labour market institutions. This data set contains observations from 1960 to 2007.

3.A.1 Economic variables

The economic variables are available at a yearly frequency in OECD (2008a)[6] and missing observations are replaced with observations from earlier data bases OECD (2002), OECD (2006) and OECD (2008b).[7]

U: Unemployment rate

The standardized unemployment rate (UNR) in Economic Outlook OECD (2008a) is used as a primary data source for the unemployment rate in the OECD countries, and missing observations are replaced by the growth rate in a corresponding time series in an earlier data base, OECD (2002). Australia, Denmark, Germany, Spain and Switzerland are prolonged by the formula in equation (A3):

$$Y_{it} = Y_{it+1} * \frac{X_{it}}{X_{it+1}} \tag{A1}$$

where Y_{it} denotes (UNR) in OECD (2008a) and X_{it} denotes the (UNR) in the earlier data base OECD (2002) for country i in time period t.

Australia and Denmark are prolonged five years backwards. Germany from 1991, Spain from 1976, Switzerland from 1969 and backwards.

3.A.2 Labour market institutions

New information for institutional variables is available every second or fifth year. Labour market institutions such as the tax wedge, union density, coordination among wage setters, and benefit replacement ratio and duration are used in this paper. The variables and the method of combining data sources are discussed in detail in the next sections.

[6]Data are collected and organized by the author. This implies that neither OECD nor any other source is responsible for the analysis or the interpretation of the data in this paper.

[7]An comprehensive overview of data and data sources is available upon request.

TW: Tax wedge

The rates described here are calculated from actual tax payments. The total tax wedge is equal to the sum of the employment tax rate (t1), the direct tax rate (t2) and the indirect tax rate (t3), as given in Equation (A4).

$$TW = t1 + t2 + t3 \tag{A2}$$

$t1$ is equal to employers' total wage costs calculated by the sum of wages received by employees and taxes payed by the employer to the government. This gives the following relationship; $t1 = SSRG/(IE - SSRG)$, where SSRG is social security contributions and IE is compensation to employees. The latter two consist of two main components, wages and salaries and social contributions. Social contributions are payed by the employers to social security schemes or private funded social insurance schemes. $t2$, are direct taxes payed by the households (TAXh) divided by current receipts of households (CRh), i.e. $t2 = TAXh/CRh$. Finally $t3 = (TAXind - SUB)/Cp$, where TAXind are net indirect taxes, SUB is the value of subsidies and Cp is the value of private final consumption expenditure.

The main data source for tax wedges is OECD (2008c) which contains information for the period 1960 to 2010. The latter years are predictions. The tax rates are calculated by the formulas above, and when a tax rate is missing, the growth rate in the same tax rate but from the data base of Nickell (2006) in the period 1960 to 2003 is used to prolong the time series for the following countries: Belgium is prolonged before 1965, Denmark is prolonged before 1966, Germany before 1970, Portugal is prolonged in the period 1960 to 1995 and Switzerland is prolonged before 1990 with the tax rates in OECD (2008c). Tax rates for Australia, Austria, Canada, Finland, France, Ireland, Italy, Japan, Netherlands, Norway, Spain, Sweden, United Kingdom and United States are not prolonged and are taken directly from the main data source OECD (2008c). New Zealand has the main data source Nickell et al. (2005) for the period 1975 to 1986 due to missing observations in OECD (2008c). Time series for $t1$, $t2$ and $t3$ from Nickell (2006) are used to extend the main data source: The growth rate of the sum of $t1$ and $t2$ is used before 1975, and the growth rate in $t3$ after 1986. Note also that the $t3$ is interpolated due to one missing observation in 1991.

BRR: Benefit replacement rates

The benefit replacement ratio is a measure of how much each unemployed worker receives in benefits from the government. The benefit replacement ratio is described in detail below.

The detailed rate for unemployment benefits divides data in three different family types: single, with a dependent spouse and with a working spouse. The benefits also depend on the employment situation: 67 percent and 100 percent of the average earnings. Within these groups, benefits are divided into the duration of benefits when being unemployed. One variable for how much each of the former groups receives in the first year, the second and third year and the fourth and fifth year. This results in six different groups: brr67a1, brr67a2, brr67a4, brr100a1, brr100a2 and brr100a4.

brr67a1: First year benefit replacement rate for workers with 67 percent of average earnings and the average over family types.

brr67a2: Benefit replacement rate for the second and third year. 67 percent of average earnings and the average over family types

brr67a4: Benefit replacement rate for the fourth and fifth year. 67 percent of average earnings and the average over family types

brr100a1, brr100a2 and brr100a4: The same as the former but for 100 percent of average earnings.

The main source for the more detailed benefit ratios is tables in employment outlook, see OECD (2004). Observations are provided every second year from 1961 to 2001. The time series are interpolated over the years, and extracted by the last known observation.

BD: Benefit duration

Benefit duration is a measure of how long the benefits last when being unemployed. The ratio is calculated by the time series described under benefit replacement rates and equation (A5).

$$BDj_{it} = \alpha \frac{brrja2_{it}}{brrja1_{it}} + (1-\alpha)\frac{brrja4_{it}}{brrja1_{it}} \tag{A3}$$

where $\alpha = 0.6$, $j = \{67, 100\}$, $i = 1, 2...20$ and $t = 1960, 1961...2007$. $brrja1_{it}$ is the benefit replacement rate in year 1, $brrja2_{it}$ is the benefit replacement rate in year 2 and 3, and finally, $brrja4_{it}$ is the benefit replacement rate in years 4 and 5. $\alpha = 0.6$ gives more weight to the second and third year as compared to the fourth and fifth year. The index is calculated for both employment situations, i.e. 67 percent and 100 percent of average earnings.

The average of $bd67_{it}$ and $bd100_{it}$ is used as an indicator of benefit duration, i.e. BD_{it}.

If benefit duration stops after one year, then brr67a2 = brr67a4 = 0, and BD67 = 0. If benefit provision is constant over the years, then brr67a1 = brr67a2 = brr67a4, and BD67 = 1. However, some countries increase payments over time and the value of benefit duration is above one.

UDNET: Union density

Union density rates are constructed using the number of union memberships divided by the number of employed.

The main data source is Visser (2009), where they have mainly calculated the trade union density index based on surveys. When data were unavailable, they have used administrative data adjusted for non-active and self-employed members. The database Nickell (2006) contains additional information for Sweden before 1975 and Ireland in 1960.

The time series for Sweden in the latter source is interpolated, and this growth rate is then used to prolong the original time series from Visser (2009).

The interaction terms between union density and coordination are prolonged by the last known observation for these countries.

CO: Coordination of wage setting

The index for coordination of wage setting describes the coordination level in the wage setting. The index ranges from 1 to 5, and the most coordinated countries have an index equal to 5:

The main source is OECD (2004), see table 3.5. The frequency for observations are five-year intervals over the period 1970-2000. The years are interpolated between means, i.e. between 1972-1977, and with 1970 and 2000 equal to the first and last five-year intervals. In the period 1960 to 1970, the observations are prolonged backwards by the last known observation for all countries. The same procedure is used to extend the time series until 2007.

EPL: Employment protection

The time series for employment protection measures the strictness of the employment protection for the employer. The overall measure for employment protection is measured on a scale from 0 to 5. Strictness is increasing in scale. Some other measures only measure the employment protection for regular- or temporary employment.

The time series for employment protection is provided by OECD (2004) for the period 1985 to 2004. The time series are based on the point observations in Annex 2.A2. 6. Note, however, that OECD (2004) claims that judgement is made when constructing the time series. That implies that time series for employment protection not are only a linear interpolation between the point observations. The measure of employment protection refers to the protection of overall employment (EPL).

Before 1985, the time series are prolonged backwards using data from Belot and van Ours (2004). The source contains data in five-year intervals, but the data are here used annually by interpolation between the means of the observation points. The percentage change is used to prolong the time series in OECD (2004), by equation (A6):

$$Y_t = Y_{t+1} * \frac{X_t}{X_{t+1}} \; where \; Y = epl \; and \; X = ERTOT\_bo \tag{A4}$$

Portugal and Spain are prolonged backwards by the last known observation in the period 1960 to 1984. The United States is prolonged backwards by the last known observation in the period 1960 to 1982.

3.B

Table B1: Estimates from the fixed effect model without dummy variable.

	All countries			Heterogenous residuals			Reduced dataset[a]		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period	1.39	0.03	0.00	1.49	0.03	0.00	1.40	0.04	0.00
Unemployment two years ago	-0.56	0.06	0.00	-0.75	0.06	0.00	-0.57	0.06	0.00
Unemployment three years ago	0.07	0.03	0.05	0.17	0.03	0.00	0.07	0.04	0.06
Employment protection (EPL), 1st difference previous period	0.08	0.29	0.78	0.06	0.25	0.81	0.10	0.31	0.76
EPL, two years ago	0.16	0.08	0.05	0.17	0.07	0.01	0.17	0.09	0.06
Benefit replacement ratio (BRR), 1st difference previous period	0.44	1.00	0.66	-0.10	0.82	0.90	0.49	1.04	0.64
BRR, two periods ago	0.95	0.29	0.00	0.72	0.23	0.00	0.94	0.32	0.00
Benefit duration (BD), 1st difference previous period	-0.87	0.67	0.20	-0.33	0.50	0.52	-0.79	0.69	0.25
BD, two periods ago	-0.19	0.21	0.37	-0.03	0.17	0.85	-0.13	0.22	0.54
Interaction - BRR and BD 1st difference previous period	1.24	2.49	0.62	0.03	1.89	0.99	1.13	2.65	0.67
Interaction - BRR and BD two periods ago	1.66	0.77	0.03	1.18	0.60	0.05	1.82	0.84	0.03
Interaction - CO and UDNET 1st difference previous period	-2.40	2.65	0.36	-0.86	2.82	0.76	-3.05	3.00	0.31
Interaction - CO and UDNET two periods ago	-1.22	0.57	0.03	-1.49	0.43	0.00	-1.53	0.65	0.02
Interaction - CO and TW 1st difference previous period	-13.29	3.04	0.00	-6.90	2.34	0.00	-14.20	3.31	0.00
Interaction - CO and TW two periods ago	-0.48	0.99	0.63	-0.33	0.78	0.67	-0.25	1.03	0.81
Union density (UDNET), 1st difference previous period	-0.22	2.49	0.93	-1.11	2.21	0.61	0.18	2.82	0.95
UDNET, two periods ago	0.66	0.35	0.06	0.49	0.32	0.12	0.85	0.38	0.03
Coordination (CO), 1st difference previous period	-0.54	0.20	0.01	-0.51	0.20	0.01	-0.47	0.22	0.03
CO, two periods ago	-0.06	0.05	0.22	-0.03	0.05	0.56	-0.04	0.05	0.40
Tax rate (TW), 1st difference previous period	0.03	1.87	0.99	1.72	1.57	0.27	0.55	1.99	0.78
TW, two periods ago	1.28	0.66	0.05	1.14	0.55	0.04	1.15	0.70	0.10
Tot. obs and the number of countries	837	20		837	20		761	18	
Standard deviation of residuals	0.7			0.7			0.7		
χ^2 of all the exogenous variables.[b]	66.78	(0.00)		51.98	(0.00)		60.92	(0.00)	
χ^2 of institutional variables (level).[b]	66.78	(0.00)		51.98	(0.00)		60.92	(0.00)	
χ^2 of institutional variables (interaction).[b]	28.63	(0.00)		25.85	(0.00)		28.49	(0.00)	
1st order autocorrelation[b]	-0.18	(0.86)		-0.18	(0.86)		0.51	(0.61)	
2nd order autocorrelation[b]	0.61	(0.54)		0.61	(0.54)		-1.48	(0.14)	

a) Without New Zealand and Switzerland.
b) Numbers in parenthesis are p-values for the relevant null.

3.C Additional results for section 3.2.

We first find the coefficients in the reduced form for the mode in equation (18). For re_t :

$$\begin{aligned}
l &= 1 - \theta_w\, \omega\, \psi_{qw}\, (1-\phi)/\chi, \\
k &= (\theta_q - \theta_w \psi_{qw})/\chi, \\
e &= 1 - (\psi_{qpi} + \psi_{qw}\, \psi_{wp}\, (1-\phi))/\chi, \quad = 0 \text{ if dynamic homogeneity} \\
n &= (\mu_q + \mu_w\, \psi_{qw})/\chi, \\
d &= (m_q\, \theta_q + c_q + (m_w\, \theta_w + c_w)\, \psi_{qw})/\chi,
\end{aligned}$$

where the denominator is: $\chi = 1 - \psi_{qw}(\phi\psi_{wp} + \psi_{wq}) > 0$. For ws_t:

$$\begin{aligned}
\lambda &= \theta_w\, \omega\, (1-\psi_{qw})(1-\phi)/\chi, \\
\kappa &= 1 - (\theta_w\, (1-\psi_{qw}) + \theta_q\, (1 - \psi_{wq} - \phi\, \psi_{wp}))/\chi, \\
\xi &= (\psi_{wp}\, (1-\psi_{qw})(1-\phi) - \psi_{qpi}\, (1 - \psi_{wq} - \phi\, \psi_{wp}))/\chi, \qquad = 0 \text{ if dynamic homogeneity} \\
\eta &= (\mu_w\, (1-\psi_{qw}) - \mu_q\, (1 - \psi_{wq} - \phi\, \psi_{wp}))/\chi, \\
\delta &= ((m_w\, \theta_w + c_w)(1-\psi_{qw}) - (m_q\, \theta_q + c_q)(1 - \psi_{wq} - \phi\, \psi_{wp}))/\chi.
\end{aligned}$$

By inspection, it is clear that all coefficient are non-negative for reasonable values of the structural coefficients. The exception is δ which can be both positive and negative. The first two disturbances in the reduced form are

$$\epsilon_{re,t} = (\varepsilon_{q,t} + \psi_{qw}\, \varepsilon_{w,t})/\chi \quad \text{and} \quad \epsilon_{prw,t} = (\varepsilon_{q,t}\, (1 - \psi_{wq} - \phi\, \psi_{wp}) - \varepsilon_{w,t}\, (1-\psi_{qw}))/\chi,$$

while the third is identical to ε_{ut} in the unemployment equation.

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Chapter 4

Do government purchases affect unemployment?

Steinar Holden and Victoria Sparrman

Abstract During the financial crisis, most OECD countries used the fiscal policy extensively to combat the crisis by stimulating the economy. More recently, fiscal policy has been reversed in many countries. The large changes in policy raise several key questions in relation to the effect on unemployment; in particular whether fiscal policy measures can be used to combat increasing unemployment, and if fiscal tightening is likely to lead to persistent high unemployment. We test the quantitative importance of government purchases for the evolution of unemployment in OECD. Compared to earlier studies we use a sample with more variation in unemployment and in institutions. We find that increased government purchases leads to lower unemployment; the point estimate is that an increase equal to one percent of GDP reduces unemployment by 0.2 percentage point in the same year, and 0.25 percentage point the year after. The effect is greater in downturns than in booms, and also greater under a fixed exchange rate regime than under a floating regime.

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4.1 Introduction

During the financial crisis, most OECD countries used the fiscal policy extensively to combat the crisis by stimulating the economy. More recently, increasing public debt and a sharp rise in default premia on sovereign debt for some countries have led to plans of substantial fiscal tightening in many countries. At the same time, unemployed has soared in the OECD area, especially in Spain, Ireland and the US. The large changes in policy and unemployment rates raise several key questions; in particular whether fiscal policy measures can be used to combat increasing unemployment, and if fiscal tightening is likely to lead to persistent high unemployment.

This paper explores the effect of fiscal policy in the form of a change in government purchases on goods and services on aggregate unemployment, using panel data for 20 OECD countries for the period 1960-2007. Our analysis builds on previous research on equilibrium unemployment by Nymoen and Sparrman (2010), who use a framework with annual panel data that draws considerably on Layard et al. (2005) and Nickell et al. (2005). In the present study we start out from the empirical unemployment relationship derived by Nymoen and Sparrman (2010), and add the change in government purchases as an explanatory variable, so as to explore the effect of fiscal policy controlling for changes in the export market and institutions in the labour market.

Whether and possibly to what extent one should use fiscal policy to stabilize the economy is a question subject to a lot of political controversy. In contrast, more concrete questions about effects should be less controversial. For example, if government purchases increase by one percent, how will this affect unemployment? - should in principle be a straightforward empirical question. Clearly, the effect is likely to vary with the circumstances and the specific policy, yet there should be a clear objective answer. However, there is no agreement in the literature on the effects of fiscal policy. To some extent, this may reflect that there has been little empirical research on this issue for decades. There is now rapid growth in the literature, and hopefully a more consensus view may emerge.

One methodological problem is that fiscal policy is potentially endogenous, in the sense that fiscal policy decisions clearly depend on the state of the economy. The large majority of the studies in the literature deal with this problem in one of three ways; see Perotti (2007), Beetsma and Giuliodori (2010), and Hall (2009) for recent reviews. Some studies focus on the effect of specific events that can be thought to be exogenous, such as changes in military spending in a response to political changes, see e.g. Ramey and Shapiro (1999) and Eichenbaum and Fisher (2005). Other studies explore the effect of fiscal policy within a structural vector autoregression (SVAR) model, where the model explains several macroeconomic variables by their lags and exogenous shocks to the variables in the model, see e.g. Blanchard and Perotti (2002), Fatás and Mihov (2001), Beetsma and Giuliodori (2010) and Monacelli et al. (2010). Finally, a number of studies analyze the effect of fiscal policy within structural models used for macro policymaking, typically of the Dynamic Stochastic General Equilibrium (DSGE) type, see e.g. Coenen et al. (2010). The various approaches have different strengths and weaknesses, which are discussed in the literature (see below).

We focus on the effect of a change in government purchases, for which we will argue that the endogeneity problem is less severe than it is for the other main budget components, taxes and transfers. Tax and transfers are given in rules and legislation, implying that changes in tax bases and the number of transfer claimants (e.g. unemployed and retirees) more or less directly lead to changes in revenues and expenditure. In contrast, government purchases are not directly linked to the state of the economy. However, as the state of the economy clearly also affects purchase decisions, we address the potential endogeneity by use of instrumental variables and by controlling for omitted variables. As compared to most existing studies, our study has the advantage of using a more extensive panel data set, which makes it possible to explore whether the effect of fiscal policy depends on a host of other factors, like the cyclical situation of the economy, the monetary regime, the openness of the economy, the type of fiscal impulse, etc. Overall, we view our study as complementary to the existing literature in this field.

We find that an increase in government purchases leads to an economically and statistically significant reduction in unemployment. The point estimate in our base equation implies that a permanent increase in government purchases equal to one percent of GDP leads to a reduction in unemployment of 0.25 percentage points after one year. The effect on unemployment decreases gradually, and only about 1/5 remains after ten years. The size of the effect is highly dependent on other factors in the economy. For example, we find that the reduction in unemployment due to a rise in government purchases is greater when the economy is in a weak cyclical situation. Furthermore, in line with the Mundell-Fleming model, we find a strong effect of fiscal policy on unemployment in countries with a fixed exchange rate, but no effect for countries with a floating exchange rate.

The rest of the paper is organized as follow. Section 4.2 provides a brief review of the theoretical and empirical literature on the effect of fiscal policy on GDP and unemployment. In section 4.3, we present our empirical approach, while the empirical results are laid out in section 4.4. Section 4.5 concludes.

4.2 The effect of fiscal policy on GDP and unemployment - theory and evidence

The aim of the paper is to explore the effect of changes in government purchases on goods and services on aggregate unemployment. However, as most of the literature focuses on the effect of GDP, and an increase in GDP usually is associated with a reduction in unemployment, we first review the effect on GDP. According to traditional Keynesian theory, the effect of fiscal policy comes via the effect on aggregate demand. Increased government spending has a direct positive effect (by being part of GDP), and indirect effects via private consumption and investment. The increase in GDP arising from the direct effect will have a positive impact on private consumption, due to the increase in contemporaneous private disposable income. However, increased government spending is also likely to lead to a higher interest rate, depending on the monetary policy regime, which will have a negative effect on private consumption and investment. Overall, GDP will increase, but the multiplier may be smaller or greater than unity.

In a neoclassical model Baxter and King (1993), the effect on output depends on the effect on the labour supply. Higher government spending will imply a negative wealth effect for households (as the spending sooner or later must be tax-financed), leading to increased labour supply and thus increased employment and output. Increased employment will raise the marginal return to capital, and thus also increase investment. However, the negative wealth effect will imply that private consumption falls. Furthermore, if higher government spending is financed by distortionary taxation, this will have a negative effect on labour supply, possibly leading to lower output and lower investment.

New Keynesian models exhibit sticky wages or prices, thus allowing for effects of aggregate demand. Nevertheless, the effects of government spending in general mimics those found in the neoclassical model. Under strict inflation targeting, the government expenditure multiplier is the same as with flexible prices. Flexible inflation targeting leads to a greater multiplier, but the multiplier of a temporary increase in government purchases is still below unity Woodford (2010). However, if there are restrictions on the monetary policy, the effects might be larger. If the real interest rate is kept constant, the multiplier is unity as there is no effect on private consumption Woodford (2010). If the real interest falls, for example because the nominal interest rate is constant (e.g. for a small country in a monetary union, or because of a zero bound restriction), while inflation increases, the government multiplier might be considerably above unity (see e.g. Christiano et al. (2009)).

Without restrictions on the monetary policy, the negative wealth effect for the households implies, as noted above, that private consumption will fall, in contrast to the traditional keynesian assertion. There have, however, been several suggestions of mechanisms that can reverse the negative wealth effect. Devereux et al. (1996) present a model where higher government spending increases the equilibrium number of firms in the intermediate good sector, where there is increasing returns to specialization. This results in a rise in productivity, leading to increased demand for labour, higher real wages and increased consumption. More recently, Galí et al. (2007) consider a model which includes rule-of-thumb consumers, that is, a share of the households consume their whole income, without any borrowing or saving. They show that the interaction of sticky prices and rule-of-thumb consumers implies that real wages increase when government consumption rise, and this also leads to an increase in private consumption.

The empirical research on the effect of fiscal policy involves a number of methodological issues. The issue that has received most attention is that government spending is likely to be endogenous, in the sense that fiscal policy decisions are affected by the state of the economy, implying that a correlation between fiscal policy and GDP or unemployment might reflect reverse causality. Thus, much of the recent research attempting to analyze the effect of fiscal policy have used methods designed to address the endogeneity issue. While the large majority of the empirical studies focus on the effect on GDP, and to some extent also the effect on private consumption, the empirical problems are to a large extent the same also for us, who study the effect on unemployment.

Some studies focus on the effect of specific events that can be thought to be exogenous, such as changes in military spending in a response to political changes, see e.g. Ramey and Shapiro (1999) and Eichenbaum and Fisher (2005). The key advantage of this approach is that with an appropriate choice of events, one avoids the endogeneity problem. However, the method also implies other weaknesses. First, there is a risk that the political changes that induce changes in military spending are also associated with other changes in the economy. As argued by Monacelli et al. (2010), it is questionable whether the effect of fiscal policy will be the same in a war period, with heavy rationing and public regulation of production, as it will be in normal periods. Private sector expectations are also likely to being strongly influenced by the political changes, in particular if there is a war, and this may also affect the evolution of the economy. Second, it is also questionable whether an increase in military spending will have the same effect on unemployment as other types of public spending, like increases employment in e.g. education.[1]

Another popular approach to study the effect of fiscal policy, is to use a structural vector autoregression (SVAR or VAR) model, where the model explains several macroeconomic variables by their lags and exogenous shocks to the variables in the model, see e.g. Blanchard and Perotti (2002), Fatás and Mihov (2001), Beetsma and Giuliodori (2010) and Monacelli et al. (2010). VAR-studies attempt to find the effects of unanticipated fiscal shocks on the economy, where the fiscal shock is identified as a deviation from a systematic fiscal policy. A crucial part of these studies is the assumptions that are made to identify the shocks, and Caldra (2010) shows that differences in results to a large extent depend on different identifying assumptions. Furthermore, as pointed out by e.g. Monacelli et al. (2010), one potential problem is that changes in government spending that appear unpredictable to the econometrician might well have been anticipated by the

[1] In a recent study, Cohen et al. (2010) show that fiscal spending shocks have a negative impact on corporate sector investment and employment. The key innovation of this study is that the spending shocks are identified by changes in congressional committee chairmanship, which is shown to have a sizeable impact on federal expenditures at state-level. However, the study does not explore the effect on GDP or unemployment.

private sector, for example from budget plans and decisions. A different problem is that one may question whether private agents in the economy are aware of the systematic fiscal rule. The systematic fiscal rule is estimated on the whole sample, while agents would only know the history up till the relevant year. Even for the recent history, it is not clear to what extent private sector agents' perceptions of the fiscal policy corresponds to the systematic rule estimated by the econometrician based on a very limited set of macroeconomic variables.

A third popular way of analyzing the effect of fiscal policy is within structural models used for macro policymaking, typically of the Dynamic Stochastic General Equilibrium (DSGE) type, see e.g. Coenen et al. (2010), or large scale estimated models Duell et al. (2009). Such studies have the great advantage that one explores the effect within a consistent theoretical framework, and where the model is estimated/calibrated also to satisfy certain empirical criteria. However, the reliability of the results clearly depend on whether the model captures the main mechanisms of the economy, and it is fair to say that also these models have their weaknesses, cf. e.g. the discussion of DSGE models in Chari et al. (2009).

Overall, our view is that the methods have their different strengths and weaknesses, which make them complementary.

As noted above, the main focus in the literature has been on the effects on GDP, private consumption and to some extent also private investment. However, there are also some studies that analyze the effect on unemployment, i.e. the topic of the current study. Holmlund and Linden (1993) explore the effects of public employment in a calibrated search model, and find ambiguous effects on unemployment as increased wage pressure may counteract the direct unemployment-reducing effect of increased public employment. More recently, Gomes (2010) shows that in a search-matching model, higher public employment leads to lower unemployment, and he also finds empirical support for this result. Monacelli et al. (2010) explore the effect of government consumption in a neoclassical model augmented with search and matching frictions. They show that while higher government consumption increases the hiring rate due to the negative wealth effect inducing higher labour supply, this effect is dampened by the rise in the real interest rate. Overall, the effect is a fairly small reduction in unemployment, and smaller than what the authors find in their empirical study, which is a structural VAR analysis on US data. In contrast, Brückner and Pappa (2010), in an analysis of 10 OECD countries using structural VARs, find that a typical estimate from the impulse responses implies that a 10 percent increase in government expenditures increases the unemployment rate at peak (which varies from $3-16$ quarters) of around $0.2-0.5$ percent. Brückner and Pappa (2010) explain the difference in results compared to Monacelli et al. (2010) as due to different sample period, arguing that the increased government spending raises unemployment after 1975. Note however that Brückner and Pappa (2010) also find that increased government spending leads to higher GDP and higher employment, so that the increase in unemployment is caused by higher participation rates due to increased labour supply.

We consider the effect of a change in government purchases on unemployment, building on a panel data estimation framework derived by Nymoen and Sparrman (2010). More specifically, we add the real change in government purchases, measured as a share of trend-GDP, to an empirical equation for aggregate unemployment as a function of a number of labour market institutions. This approach has several advantages. First, an extensive literature has shown that aggregate unemployment to a large extent is determined by labour market institutions, see e.g. Layard et al. (2005) and Nickell et al. (2005). Thus, it seems expedient to control for the effect of labour market variables when analysing the effect of fiscal policy. Second, with a data set covering 20 countries and 47 years, there is large variation in a number of key variables, making it possible to explore how the effect of fiscal policy may vary depending on for instance the monetary regime, the openness of the economy, or the size of the public debt.

As noted in the introduction, the fact that we restrict attention to the effect of government purchases makes the endogeneity problems less compelling than if we had analysed the effect of changes in taxes and transfers. Tax revenues and expenditures on transfers are clearly endogenous, following changes in the economy according to rules and legislation. In contrast, government purchases are not directly linked to the state of the economy. For example, all "passive" unemployment expenditure like benefits, and the large majority of active unemployment related expenditure are classified as transfers, not government purchases, and thus not included in our analysis. Clearly, the state of the economy also affects purchase decisions, but also other factors come into play, like electoral cycles, party politics, lobbyism and pressure groups, media attention, etc. Furthermore, a large part of government purchases may be subject to a lengthy bureaucratic process involving both the decision making and the implementation, implying that there is no clear cut or simple relationship between the state of the economy and government purchases. Yet there is of course a potential endogeneity problem: When we include the change in government purchases as a regressor in the estimated equation, any correlation between our fiscal variable and the error term in the equation may bias our estimate.

We address this problem in two ways. First, we use instrumental variable estimation, where we treat the measure of fiscal policy as endogenous. However, it is hard to find appropriate instruments, even if the ones we use (past values of the change in government purchases and past values of government debt) seem to pass the most basic tests.

Second, we use an omitted variables approach. The idea here is that fiscal policy might be correlated with the error term because it is affected by other explanatory variables that also affect unemployment. For example, fiscal policy might be pro-cyclical, because in a boom, tax revenues increase making it possible for politicians to increase public spending. At the same time, the increase in tax revenues during the boom might be correlated with a fall in unemployment. However, in this case including tax revenues as a regressor in the unemployment equation would lend fiscal policy uncorrelated with the error term, removing the bias in the coefficient. Likewise, fiscal policy might be countercyclical because the government wants to stabilize the economy, and thus increase spending whenever GDP growth falls. As a fall in GDP growth also will be associated with a rise in unemployment, this policy will involve a correlation between fiscal policy and the error term. However, by including GDP growth as a regressor in the unemployment equation, the correlation between fiscal policy and the error term will be removed, and there is no bias in the coefficient for fiscal policy. Admittedly, it is impossible to include all possible omitted variables, so one would not expect this approach to remove any correlation between fiscal policy and the error term. However, by including the most likely omitted variables, one should be able to reduce any possible bias substantially.

From a theoretical perspective, one would expect the effect of an increase in government spending to depend on whether it is deficit-financed or debt-financed. However, to distinguish between these two alternatives one needs a measure of discretionary changes in taxes, so that one can circumvent the problem that an upswing in the economy at the same time will involve lower unemployment and higher tax revenues. However, as noted above, we view the endogeneity problem as larger for taxes than for government purchases, and at the present stage we have chosen not to include a measure of discretionary tax changes. Because of this, we are unable to distinguish between deficit-financed and debt-financed government spending. Thus, our results must be interpreted as an average effect, where the weights depend on the average method of financing over the sample period. As government debt has increased in most countries over the sample period, the increase in government spending is partly debt-financed and partly deficit-financed.

4.3 Empirical Approach

There is now a vast literature arguing that unemployment is affected by labour market institutions; see e.g. Layard et al. (2005) and Nickell et al. (2005). Changes in labour market variables are also important for explaining the evolution of unemployment over time, cf. Nickell et al. (2005). As we are interested in the effect of government purchases, we want to control for other effects by use of the unemployment equation derived in Nymoen and Sparrman (2010), which in accordance with Layard et al. (2005) and Nickell et al. (2005), is a function of labour market institutions and shocks. Specifically, Nymoen and Sparrman (2010) derive a preferred equation on the form

$$u_{it} = \beta_{0i} + \beta_1 u_{it-1} + \beta_2 u_{it-2} + \beta_3 u_{it-3} + \beta_4 \mathcal{I}_{it-1} + \beta_5 \mathcal{I}_{it-2} + \beta_6 \mathcal{D}_{it} + \epsilon_{it}$$

where u_{it} is unemployment in country i in period t, $\mathcal{I}_{it-1}$ is a vector of institutional variables, and $\mathcal{D}_{it}$ is a vector of structural breaks (detected by a large outliers approach), capturing other important shocks that might affect unemployment. The parameters are functions of the underlying model; for example the theory implies that $\beta_1 > 0$ and $\beta_2 < 0$. The dynamic structure follows from the theoretical labour market framework of Nymoen and Sparrman (2010), where the system of wage and price curves implies that the autoregressive order of the unemployment equation should at least be three.

To this equation we add the real change in the government purchases on goods and services, measured as a share of trend GDP. This variable (dG) is calculated as the real growth rate of government purchases (i.e. real government consumption and investments, not including government transfers), multiplied by government purchases as a share of trend of GDP, in nominal prices, see appendix 4.A for details and calculations.

The decade averages of the change in government purchases are shown in table 1. We observe that government purchases have generally increased in real terms, and more so in the 1960s and 70s. There is however considerable variation within and across countries.

The motivation behind this specification is to ensure that our left hand side variable only measures changes in government purchases, and that it not directly affected by a change in GDP. Some studies (e.g. Alesina and Ardagna (2009), Duell et al. (2009) consider the effect of a change in the ratio of government spending to GDP. This choice involves the risk that a reduction in GDP caused by e.g. an external shock leads to an increase in the ratio of government spending to GDP, which would then be associated with a reduction in GDP. For this reason we also use a backward-looking measure of trend GDP, where the trend real growth is measured as the moving average of the growth rate over the past ten years. With a two-sided measure of trend-growth, there would be a risk that the future evolution of GDP affects the estimated trend-GDP, implying a possibility that the future evolution of GDP affects the measure of contemporaneous fiscal policy. We have also explored the effect of fiscal indicator of Braconier and Holden (1999), which is similar to the change in the ratio of government purchases to trend GDP, with essentially the same results.

We also include another shock variable, which is an indicator for the cyclical state of the economy of the trading partners. More specifically, the indicator is calculated as a weighted average of the GDP-gap of the trading partners, where the GDP-gap is the deviation of GDP from Hodrick Prescott-trend, divided by the trend, and the weights reflect the share of the exports from country i that goes to each of the trading partners j. The motivation for including this variable, which we term export market, is the idea that the cyclical situation among the trading partners affects the demand for a country's export, which again may affect GDP, unemployment and fiscal policy in the country. Including the export market indicator may improve the precision of the estimates, as well as reduce any possible bias in our coefficient of interest, to the extent that the added explanatory variable is correlated with the change in fiscal policy.

Table 1: Growth in government purchases - country specific mean and standard deviation

Country stats	**1960-69**	**1970-79**	**1980-89**	**1990-99**	**2000-07**	**1960-07**
Australia mean	1.38	0.87	0.87	0.64	0.80	0.90
sd	0.75	0.60	0.57	0.26	0.21	0.55
Austria mean	0.93	0.89	0.23	0.48	0.13	0.53
sd	0.48	0.35	0.24	0.29	0.30	0.46
Belgium mean	1.55	1.36	0.17	0.35	0.43	0.75
sd	0.52	0.40	0.53	0.37	0.29	0.70
Canada mean	1.56	1.16	0.72	0.25	0.82	0.88
sd	0.59	0.70	0.36	0.52	0.20	0.66
Denmark mean	.	0.89	0.16	0.54	0.52	0.51
sd	.	0.69	0.52	0.45	0.31	0.55
Finland mean	1.31	1.19	0.79	0.20	0.37	0.77
sd	1.07	0.72	0.32	0.85	0.35	0.81
France mean	1.13	0.89	0.80	0.44	0.49	0.72
sd	0.21	0.35	0.24	0.41	0.20	0.38
Germany mean	1.47	1.22	0.19	0.45	0.13	0.68
sd	0.94	0.58	0.51	0.48	0.21	0.78
Ireland mean	1.06	1.59	-0.04	1.11	1.45	1.02
sd	0.57	0.81	1.41	0.50	1.37	1.13
Italy mean	0.88	0.70	0.59	0.04	0.40	0.51
sd	0.15	0.19	0.24	0.37	0.17	0.37
Japan mean	2.17	1.79	0.53	0.73	-0.08	1.03
sd	0.85	1.44	0.47	0.62	0.21	1.14
Netherlands mean	1.12	0.78	0.67	0.65	0.93	0.81
sd	0.37	0.76	0.32	0.24	0.76	0.54
New Zealand mean	0.73	0.97	0.28	0.54	0.93	0.68
sd	1.07	1.49	0.94	1.01	1.14	1.13
Norway mean	1.64	1.56	0.87	1.18	0.90	1.22
sd	0.58	0.45	0.53	0.61	0.53	0.61
Portugal mean	.	1.80	0.93	0.94	0.21	0.80
sd	.	0.83	0.66	0.61	0.61	0.74
Spain mean	0.49	0.99	0.96	0.79	1.18	0.91
sd	0.18	0.27	0.50	0.46	0.18	0.40
Sweden mean	1.86	1.10	0.50	0.41	0.24	0.80
sd	0.49	0.70	0.31	0.56	0.45	0.76
Switzerland mean	0.37	0.22	0.36	0.10	0.07	0.21
sd	0.13	0.32	0.27	0.24	0.15	0.27
United Kingdom mean	0.70	0.53	0.20	0.30	0.74	0.46
sd	0.75	0.62	0.32	0.39	0.82	0.60
United States mean	1.20	0.12	0.67	0.26	0.44	0.51
sd	0.82	0.43	0.42	0.28	0.23	0.58
Total mean	1.25	1.00	0.52	0.52	0.56	0.74
sd	0.77	0.80	0.61	0.58	0.65	0.73

The large outliers breaks are likely to capture all types of shocks, including shocks to the fiscal policy and the export market. Thus, if we were to retain these breaks, one would expect that this would involve a downward bias in the effect of the fiscal policy. Thus, in our main estimations, we omit these breaks and estimate an equation of the following form

$$\begin{aligned} u_{it} =& \beta_{0i} + \beta_1 u_{it-1} + \beta_2 u_{it-2} + \beta_3 u_{it-3} + \beta_4 \Delta\mathcal{I}_{it-1} + \beta_5 \mathcal{I}_{it-2} \\ &+ \beta_6 \Delta dG_t + \beta_7 dG_{t-1} + \beta_8 \Delta dG_{t-1} + \beta_9 \Delta XM_t + \beta_{10} XM_{t-1} + \beta_{11} \Delta XM_{t-1} + \epsilon_{it} \end{aligned} \tag{1}$$

The evolution of the change in government purchases and the rate of unemployment are illustrated in figure 1 and 2. There may seem to be a negative comovement between these two variables in some countries like Belgium, Canada, Denmark and New Zealand, but not in others.

In most of the analysis, we use a Fixed Effects (FE) estimator, allowing for permanent country-specific differences in unemployment that are not accounted for by the other explanatory variables. A random effect model would require that there is no correlation between the country fixed effects and the explanatory variables in the model. However, this assumption is rejected in a Hausman test, although only with a p-value of 6 percent. In principle, the FE estimator is biased when the regression includes a lagged endogenous variable. However, with a long time dimension of more than 40 years, this bias is small, cf. Judson and Owen (1999). In addition, other estimations methods which avoid the sample bias also have their difficulties, cf. Roodman (2009). Some alternative estimation methods on this data set are discussed in Nymoen and Sparrman (2010).

The model is estimated on annual frequency. The main reason for this is that the unemployment equation that forms the basis for the analysis is performed on annual data. Moreover, annual data has the advantage of a allowing a much longer time span, as very few countries have quarterly data for the fiscal policy from the 1960s and 70s. Finally, it has also been argued that annual data to a better extent captures the actual fiscal decisions, as the fiscal impulses are likely to follow annual budgets, see discussion in Beetsma and Giuliodori (2010).

4.4 Empirical results

For comparison, the first model in table 2 is a reestimation of the unemployment equation in Nymoen and Sparrman (2010) where a structural break variable (detected by large outliers) is included. The structural break variable has a significant effect on unemployment, and the effect is longlasting. To focus on novel variables, the coefficients for the labour

Figure 1: The change in government purchases, subtracted country specific mean, the output gap and the unemployment rate during the period 1960 to 2007

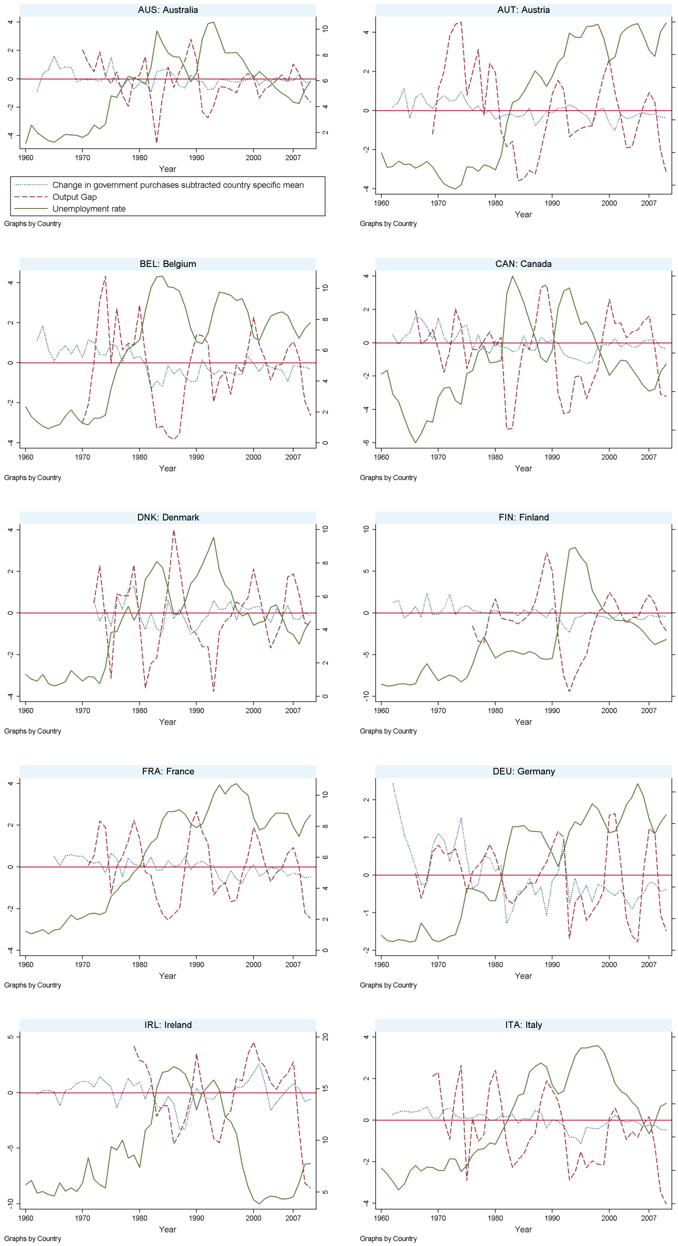


Figure 2: The change in government purchases, subtracted country specific mean, the output gap and the unemployment rate during the period 1960 to 2007

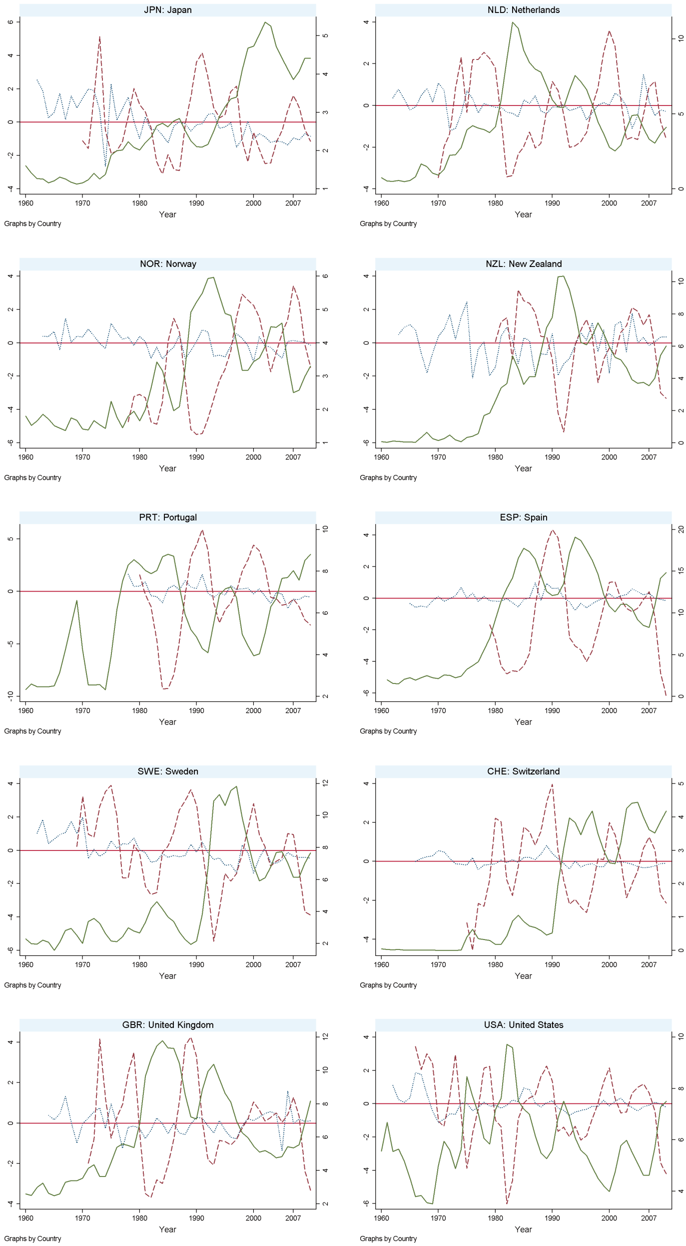


Table 2: Equation (1) with the growth in government purchases and the demand for export products

	Model 1			Model 2			Model 3		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period, (U_{t-1})	1.38	0.03	0.00	1.29	0.03	0.00	1.28	0.04	0.00
Unemployment two years ago (U_{t-2})	-0.52	0.05	0.00	-0.41	0.05	0.00	-0.39	0.05	0.00
Unemployment three years ago (U_{t-3})	0.06	0.03	0.04	0.01	0.03	0.70	-0.01	0.03	0.82
Large outlier detection[a]	0.94	0.05	0.00	0.82	0.04	0.00			
Demand components:									
Export market, 1st diff. (ΔXM_t)				-0.39	0.05	0.00	-0.53	0.06	0.00
Export market, prev. period (XM_{t-1})				0.12	0.06	0.03	0.17	0.07	0.01
Export market, 1st diff. prev. period (ΔXM_{t-1})				-0.23	0.06	0.00	-0.30	0.07	0.00
Change govt. purchases, 1st diff. (ΔdG_t)				-0.14	0.03	0.00			
Change govt. purchases, prev. period (dG_{t-1})				-0.21	0.05	0.00			
Change govt. purchases, 1st diff. prev. period (ΔdG_{t-1})				0.03	0.03	0.31			
Change govt. purchases (dG_t)							-0.19	0.04	0.00
Obs = Country*Average groups	837	20	41.9	794	20	39.7	801	20	40.0
Standard deviation of residuals	0.58			0.54			0.65		
χ^2 of all the exogenous variables.[b]	335.53	(0.00)		765.02	(0.00)		287.57	(0.00)	
χ^2 of dummy, fiscal policy and exports.[b]				635.72	(0.00)				
χ^2 of policy and exports.[b]				118.01	(0.00)		195.35	(0.00)	
1st order autocorrelation[b]	0.37	(0.71)		0.29	(0.77)		0.65	(0.52)	
2nd order autocorrelation[b]	-1.71	(0.09)		-2.00	(0.04)		0.15	(0.88)	

Estimation method: Fixed effect coeffcients estimate, standard errors from GLS (xtgls without options).
In all equations it is also controlled for labour market institutions, cf. complete table in appendix 4.B.
a) Dummy by impulse saturation.
b) Numbers in parenthesis are p-values for the relevant null.

market institutions are not included in the table. However, the complete results are found in table B1 in appendix 4.B.

In model 2, we add the change in government purchases and the export market variable, both as first differences and lagged level (for the government purchases, this is first difference of the change, and the lagged change). We observe that the government purchase variables are highly significant, and the same is true for two of the export market variables. However, we would expect the structural break to capture some of the effect of the demand variables, reducing the size of their coefficients. Thus, in model 3 we omit the structural break variable. In this specification the first difference and the lagged change in government purchases obtain essentially the same coefficient, so they can be put together, while the lagged first difference is not statistically significant. Thus, we end up with a highly significant change in government purchases, as well as three variants of the export market market which are all statistically significant. Note that result in model 3 is robust to exclusion order, i.e. the lagged first difference of the change in government purchases is not significant even if we first exclude the lag of the export market. The effect of government purchases is unaffected by including year dummies, cf. model 3 in table B1 in appendix 4.B. The coefficient values for the export market variables are however much smaller with year dummies, which suggests that these dummies may capture common shocks that affect most or all OECD countries.

The autocorrelation test in the lower part of table 2, model 3, shows no sign of autocorrelation. Figure 3 shows the estimated residuals of model 3 in table 2. Figure 3 also supports the picture of no autocorrelation, even though there is some variation across countries.

Figure 3: Estimated residuals of model 3 in table 2

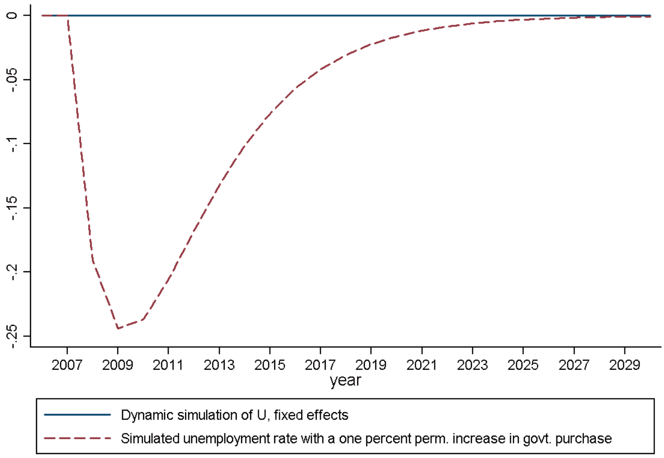


Figure 4: The effect of a permanent increase in government purchases, equal to one percent of GDP, from 2008, based on simulation of equation (1) with estimated coefficients from model 3 in table 2

Figure 4 shows the effect on unemployment over time from a permanent increase in government purchases equal to one percent of trend GDP. The maximum impact of -0.25 percentage points is reached in the second year, then the effect weakens gradually to be almost negligible after 10 years. This effect is very close to the findings in IMF (2010), based on a study of fiscal consolidations in 15 OECD countries over the last 30 years. They find that spending-based deficit cuts equal to one percent of GDP raise the unemployment rate of about 0.2 percentage points. Monacelli et al. (2010) find a larger effect on US data; an increase in government spending equal to one percent of GDP leads to a fall in the rate of unemployment of 0.6 percentage points after ten quarters. However, as noted above, Brückner and Pappa (2010) find in an analysis of 10 OECD countries using structural VARs, that a typical estimate from the impulse responses implies that a 10 percent increase in government expenditures increases the unemployment rate at peak (which varies from $3 - 16$ quarters) of around $0.2 - 0.5$ percent.

4.4.1 Robustness checks for government purchases being endogenous

As discussed above, the estimated coefficient of government purchases will be biased if government purchases also react to changes in the state of the economy that are correlated with the rate of unemployment. We deal with this problem in two ways; by using instruments for government purchases possibly being endogenous, and by controlling for omitted variables.

Table 3: Correlation of government purchases with other variables

	Change govt. purchases (dG_t)		
	Correlation	p-value	Observations
Change govt. purchases, 1st diff. prev. period (ΔdG_{t-1})	0.13	0.00	880.00
Change govt. purchases, 1st diff. prev. period (ΔdG_{t-2})	0.02	0.60	860.00
Change govt. purchases, prev. period (dG_{t-1})	0.48	0.00	900.00
Change govt. purchases, two periods ago (dG_{t-2})	0.32	0.00	880.00
Change govt. purchases, three periods ago (dG_{t-3})	0.29	0.00	860.00
Change govt. purchases, four periods ago (dG_{t-4})	0.25	0.00	840.00
Debt prev. period ($debt_{t-1}$)	-0.30	0.00	695.00
Debt two periods ago ($debt_{t-2}$)	-0.25	0.00	677.00
Debt, 1st diff. prev. period ($\Delta debt_{t-1}$)	-0.26	0.00	677.00
Debt, 1st diff. two periods ago ($\Delta debt_{t-2}$)	-0.26	0.00	657.00
N	1020		

Instrument variable approach

One way to deal with the possible endogeneity of government purchases is to find instruments that are uncorrelated with the error term and highly correlated with the change in government purchases. We use the lagged change in government purchases as well as the lagged change in public debt as a ratio to GDP. Note that while our results imply that the lagged change in government purchases is correlated with lagged unemployment, the fact that lagged unemployment is also included in the equation implies that lagged government purchases may well be a valid instrument. We have also tried election year, based on the idea that governments may pursue an expansionary fiscal policy in connection with elections to increase the probability of reelection; see evidence in Shi and Svensson (2006) (2006). However, including election year did not affect the result, and as election year is potentially endogenous in countries where the government can choose the time of the election, we decided to leave it out in the presented specification.

Table 3 shows the correlation between the change in government purchases and several specifications of the instrumental variables. We observe that several of the specifications are significantly correlated with the change in government purchases.

Table 4 shows the results of the instrumental variable estimation (for comparison, the first column displays the FE estimation from table 2). The point estimate indicates that an increase in government purchases equal to one percent of GDP reduces unemployment by half a percentage point, i.e. more than twice the effect from the FE estimates. The effect is also highly statistically significant. This suggests that government purchases is endogenous, leading to a downward bias in the coefficient in the FE result. Note however that the F-test of the instrument variables is equal to 9.2, which is at the borderline to a sign of weak instruments. Because of the difficulty of obtaining satisfying instruments, most of the subsequent regressions are ordinary FE estimates, not allowing for endogeneity, and this suggests some caution in the interpretation of the results.

Table 4: Equation (1) with instrumental variables

	FE[a]			IV[b]		
	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period	1.28	0.04	0.00	1.27	0.05	0.00
Unemployment two years ago	-0.39	0.05	0.00	-0.43	0.07	0.00
Unemployment three years ago	-0.01	0.03	0.82	0.01	0.04	0.81
Demand components:						
Export market, 1st diff. (ΔXM_t)	-0.53	0.06	0.00	-0.51	0.07	0.00
Export market, prev. period (XM_{t-1})	0.17	0.07	0.01	0.22	0.08	0.01
Export market, 1st diff. prev. period (ΔXM_{t-1})	-0.30	0.07	0.00	-0.28	0.08	0.00
Change govt. purchases (dG_t)	-0.19	0.04	0.00	-0.50	0.18	0.01
Obs = Country*Average groups	801	20	40.0	626	20	31.3
Standard deviation of residuals	0.65			0.92		
χ^2 of all the exogenous variables.[c]	287.57	(0.00)		210.14	(0.00)	
χ^2 of dummy, fiscal policy and exports.[c]	195.35	(0.00)		128.26	(0.00)	

In all equations it is also controlled for labour market institutions.

a) Estimation method: Fixed effect coefficients estimates, standard errors from GLS (xtgls without options).

b) Change govt. purchases (dG_t) is treated as endogenous. Instruments are: ΔdG_{t-1}, dG_{t-2} and $\Delta debt_{t-1}$.

c) Numbers in parenthesis are p-values for the relevant null.

Controlling for omitted variables

The possible weakness of the instruments suggests that we also address the potential endogeneity of government purchases by other means. The idea behind controlling for omitted variables is that fiscal policy might be correlated with the error term because it is affected by other explanatory variables that also are correlated with unemployment. For example, fiscal policy might be pro-cyclical, because in a boom, tax revenues increase making it possible for politicians to spend more money; this effect is termed the voracity effect by Tornell and Lane (1999). At the same time, the increase in tax revenues during the boom might be correlated with a fall in unemployment. However, in this case including tax revenues as a regressor in the unemployment equation would lend fiscal policy uncorrelated with the error term, removing the bias in the coefficient. As the government purchases are typically decided in the budget process in the fall the year prior to the budget year, it would be tax revenues for the year when the budget is decided that might affect the budget. Thus, in table 5 we include the lagged change in tax revenues as a share of trend GDP to capture that higher revenues might lead to increased government purchases. In contrast, if the government attempts to use fiscal policy to stabilize the economy, one would expect an increase in government purchases in downturns, when GDP growth is low, or the output gap is negative. To control for this, we also include GDP growth and the change in the output gap, both lagged, in table 5.

We observe that the effect of government purchases is not affected by including the additional explanatory variables in model 2 in table 5. [2] This lends considerable support

[2]The sample size is reduced somewhat because of data availability, and the isolated effect of this - i.e. model 1 on the reduced sample size - yields a coefficient for government purchases of 0.27. Thus, the isolated effect of the additional variables is a small reduction in the value of the coefficient.

to the robustness of this effect, as both the lagged GDP growth and the output gap are variables that are strongly correlated with unemployment. Note however that some of the export market variables are no longer significant. This emphasizes that including lagged GDP growth and lagged output gap entail a strong test of the explanatory power of the variables.

In model 3 and 4 in table 5, we control for the possible endogeneity of government purchases in a somewhat different way, by also including consensus forecast for GDP growth, unemployment and the output gap. Again, one might conjecture that government purchases would respond to such forecasts, and that the correlation we find between government purchases and unemployment is due to both variables being correlated with the forecasts. However, we see that the change in government purchases has a significant negative impact on unemployment even when controlling for forecasts. The coefficient value is slightly smaller, but a comparison with model 5 shows that this difference is due to the much smaller sample size in the regressions with forecasts.

Table 5: Equation (1) with control for omitted variables

	Model 1			Model 2			Model 3			Model 4			Model 5		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period	1.28	0.04	0.00	1.13	0.05	0.00	1.06	0.07	0.00	1.19	0.07	0.00	1.05	0.07	0.00
Unemployment two years ago	-0.39	0.05	0.00	-0.30	0.07	0.00	-0.34	0.09	0.00	-0.53	0.09	0.00	-0.40	0.09	0.00
Unemployment three years ago	-0.01	0.03	0.82	0.01	0.04	0.83	0.08	0.05	0.13	0.12	0.06	0.03	0.13	0.05	0.01
Demand components:															
Export market, 1st diff. (ΔXM_t)	-0.53	0.06	0.00	-0.55	0.07	0.00	-0.39	0.08	0.00	-0.42	0.09	0.00	-0.40	0.08	0.00
Export market, prev. period (XM_{t-1})	0.17	0.07	0.01	0.09	0.07	0.23	0.01	0.13	0.93	-0.08	0.14	0.56	-0.05	0.12	0.69
Export market, 1st diff. prev. period (ΔXM_{t-1})	-0.30	0.07	0.00	-0.08	0.08	0.29	0.06	0.13	0.64	0.02	0.13	0.86	0.11	0.11	0.33
Change govt. purchases, (dG_t)	-0.19	0.04	0.00	-0.21	0.05	0.00	-0.17	0.06	0.00	-0.16	0.06	0.01	-0.16	0.06	0.00
Controls:															
Log GDP 1st diff. prev. period				-20.20	4.46	0.00	-22.96	10.42	0.03				-17.58	9.08	0.05
Outputgap 1st diff. prev. period				0.06	0.05	0.22	0.04	0.10	0.68				0.01	0.09	0.95
Direct and indirect taxes divided by trend GDP, 1st diff. prev. period				1.37	3.55	0.70	-1.83	3.16	0.56				-1.47	3.19	0.64
Direct and indirect taxes divided by trend GDP, 1st diff. two periods ago				0.12	3.35	0.97	-6.76	3.14	0.03				-5.83	3.11	0.06
Year$_{t-1}$ forecast of GDP growth year t							-0.00	0.06	0.98	-0.06	0.06	0.31			
Year$_{t-1}$ forecast of output gap year t							-0.05	0.10	0.61	-0.15	0.10	0.12			
Year$_{t-1}$ forecast of unemployment year t							-0.15	0.09	0.08	-0.14	0.09	0.15			
Year$_{t-1}$ forecast of GDP growth year $t+1$							-0.02	0.08	0.85	0.00	0.09	0.98			
Year$_{t-1}$ forecast of output gap year $t+1$							0.11	0.09	0.23	0.13	0.10	0.17			
Year$_{t-1}$ forecast of unemployment year $t+1$							0.15	0.09	0.10	0.15	0.10	0.14			
Obs = Country*Average groups	801	20	40.0	616	20	30.8	207	20	10.4	207	20	10.4	207	20	10.4
Standard deviation of residuals	0.65			0.61			0.34			0.36			0.34		
χ^2 of all the exogenous variables.[a]	287.57	(0.00)		235.53	(0.00)		95.47	(0.00)		78.58	(0.00)		92.72	(0.00)	
χ^2 of policy and exports.[a]	195.35	(0.00)		139.26	(0.00)		45.20	(0.00)		41.60	(0.00)		44.72	(0.00)	
1st order autocorrelation[a]	0.65	(0.52)		1.97	(0.05)		-3.02	(0.00)		-3.00	(0.00)		-0.26	(0.79)	
2nd order autocorrelation[a]	0.15	(0.88)		-0.65	(0.52)		0.66	(0.51)		0.18	(0.86)		-1.59	(0.11)	

Estimation method: Fixed effect coefficients estimates, standard errors from GLS (xtgls without options).

In all equations it is also controlled for labour market institutions.

b) Numbers in parenthesis are p-values for the relevant null.

4.4.2 Heterogenous effects of government purchases across countries

The effect of government purchases might differ across countries, and this is explored in table 6. Model 1 presents the base specification allowing for country-specific effects of government purchases. We observe that there is considerable variation, yet the sign is negative for 17 of 20 countries, and the effect is statistically significant at the 10 percent level for 8 of 20 countries. If we include time dummies, cf. model $1i$, the effect is significant for 11 countries.

Model 2 presents the results using IV estimation. At the ten percent level, the effect of the change in government purchases is only significantly negative for two countries (Finland and France). In addition, two countries have negative coefficients with p-value of 12 and 11 percent (Ireland and Italy). Overall, the coefficient is negative for 16 of 20 countries, but generally the standard errors are large. Presumably, the results reflects the problem of weak instruments, which is clearly more important when we allow for country-specific coefficients for the effect of government purchases.

In model 3 we address the endogeneity by including possible omitted variables, and include the lagged growth rate of GDP, the lagged output gap (first difference), and indirect and direct taxes as a share of trend GDP (lagged and first differences). We find a significant negative effect of government purchases on unemployment in 8 countries, and the sign is negative in 16 of the 20 countries.

Table 6: Equation (1) with effect of government purchases cross countries

	Model 1			IV 2[a]			Model 3			Model 1i		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period	1.25	0.03	0.00	1.25	0.07	0.00	1.07	0.05	0.00	1.23	0.03	0.00
Unemployment two years ago	-0.38	0.05	0.00	-0.35	0.12	0.00	-0.24	0.07	0.00	-0.37	0.05	0.00
Unemployment three years ago	-0.00	0.03	0.90	-0.05	0.07	0.49	0.01	0.04	0.86	-0.00	0.03	0.96
Export market indicator:												
Export market, 1st diff. (ΔXM_t)	-0.52	0.06	0.00	-0.44	0.10	0.00	-0.54	0.06	0.00	-0.24	0.10	0.02
Export market, prev. period (XM_{t-1})	0.17	0.07	0.01	0.26	0.14	0.06	0.07	0.07	0.33	0.09	0.10	0.39
Export market, 1st diff. prev. period (ΔXM_{t-1})	-0.31	0.07	0.00	-0.29	0.13	0.02	-0.10	0.08	0.18	-0.29	0.10	0.00
Change govt. purchases (dG_t):												
dG_t (Australia)	-0.00	0.21	1.00	-1.15	2.26	0.61	0.10	0.21	0.64	-0.01	0.20	0.97
dG_t (Austria)	-0.19	0.28	0.51	-0.73	0.84	0.39	-0.12	0.28	0.67	-0.30	0.27	0.26
dG_t (Belgium)	-0.34	0.16	0.03	-0.53	0.44	0.23	-0.23	0.19	0.21	-0.24	0.15	0.11
dG_t (Canada)	0.10	0.16	0.54	0.18	0.36	0.61	0.33	0.17	0.05	0.19	0.15	0.20
dG_t (Denmark)	-0.35	0.20	0.08	2.51	1.66	0.13	-0.24	0.19	0.19	-0.34	0.19	0.07
dG_t (Finland)	-0.55	0.13	0.00	-1.02	0.55	0.07	-1.27	0.19	0.00	-0.59	0.13	0.00
dG_t (France)	-0.35	0.28	0.21	-1.92	1.05	0.07	-0.23	0.28	0.42	-0.42	0.26	0.10
dG_t (Germany)	-0.34	0.18	0.06	-0.65	0.51	0.20	-1.07	0.71	0.13	-0.34	0.17	0.05
dG_t (Ireland)	-0.32	0.09	0.00	-0.37	0.23	0.11	-0.25	0.10	0.01	-0.32	0.09	0.00
dG_t (Italy)	-0.76	0.29	0.01	-1.12	0.72	0.12	-0.72	0.30	0.02	-0.69	0.27	0.01
dG_t (Japan)	-0.04	0.10	0.66	-0.52	0.52	0.32	-0.05	0.11	0.63	-0.07	0.09	0.44
dG_t (Netherlands)	-0.29	0.18	0.11	-0.33	0.49	0.50	-0.38	0.19	0.05	-0.32	0.17	0.07
dG_t (Norway)	-0.37	0.17	0.03	-1.11	0.96	0.25	-0.39	0.20	0.05	-0.31	0.16	0.06
dG_t (New Zealand)	-0.04	0.10	0.71	-0.30	0.58	0.60	-0.17	0.15	0.27	-0.06	0.10	0.54
dG_t (Portugal)	-0.23	0.19	0.22	-0.94	1.00	0.35	-0.26	0.20	0.20	-0.35	0.18	0.05
dG_t (Spain)	-0.92	0.29	0.00	-0.48	0.62	0.44	-1.01	0.27	0.00	-0.98	0.27	0.00
dG_t (Sweden)	-0.15	0.16	0.35	-0.18	1.15	0.88	-0.65	0.21	0.00	-0.16	0.15	0.27
dG_t (Switzerland)	0.11	0.38	0.77	24.53	23.37	0.29	-0.21	1.06	0.84	-0.03	0.36	0.92
dG_t (United Kingdom)	0.04	0.16	0.79	-0.20	1.30	0.88	0.07	0.17	0.70	0.10	0.16	0.52
dG_t (United States)	-0.02	0.17	0.90	0.03	0.33	0.94	0.21	0.23	0.36	0.02	0.16	0.90
Controls:												
Log GDP 1st diff. prev. period							-23.29	4.52	0.00			
Outputgap 1st diff. prev. period							0.09	0.05	0.05			
Direct and indirect taxes divided by trend GDP, 1st diff. prev. period							1.24	3.43	0.72			
Direct and indirect taxes divided by trend GDP, 1st diff. two periods ago							3.01	3.26	0.36			
Obs = Country*Average groups	801	20	40.0	626	20	31.3	616	20	30.8	801	20	40.0

Estimation method: Fixed effect coefficients estimates in all models except for in the second column where IV is used, standard errors from GLS (xtgls without options).

In all equations it is also controlled for labour market institutions.

a) Change govt. purchases (dG_t) is treated as endogenous. Instruments are: ΔdG_{t-1}, dG_{t-2} and $\Delta debt_{t-1}$.

4.4.3 Does the effect of government purchases vary over the cycle?

An important question from a policy perspective is whether the effect of government purchases varies over the business cycle. We measure the cyclical situation of the economy by use of the output gap as measured by the OECD, and defined as real GDP minus trend GDP, divided by trend GDP, and multiplied by 100.

The first model in table 7, we extend equation (1) by including the interaction between the output gap and the change in government purchases. The interaction term is strongly significant, with positive sign, implying that an increase in government purchases leads to a larger reduction in unemployment in bad times when the output gap is negative than in good times. The effect of the interaction term is quite large, with estimated coefficient equal to 0.07. This means that if the output gap is positive and equal to 2 percent, an increase in government purchases equal to one percent of GDP will lower the unemployment rate by only $0.21 - 0.14 = 0.07$ percentage points at impact. In contrast, if the output gap is negative and equal to -2 percentage point, the same increase in government purchases will decrease unemployment by $0.21 + 0.14 = 0.35$ percentage points at impact.

In model 2 in table 7, we interact the interaction variable with a dummy for the output gap being positive ("boom", to explore possible asymmetries in the link between output gap and the effect of government purchases. Allowing for asymmetry results in a more complex picture. The estimated effect of a rise in government purchases increases rather strongly with the absolute value of the output gap when the output gap is negative: If the output gap is -4 percentage points, an increase in government purchases equal to one percent of GDP is estimated to reduce unemployment by $4*0.14 = 0.56$ percentage point. However, in a boom, the size of the output gap is less important, and surprisingly with the same sign, implying that the effect of a rise in government purchases is increasing in the output gap. If the output gap is $+3$ percentage points, the effect a rise in government purchases equal to one percent of GDP is a decrease in unemployment by $3*(-0.14+0.18) = 0.12$.

The third model in table 7 shows that there is no sign of non-linear effects of the output gap.

Table 7: Equation (1) with interaction with the output gap

	Model 1			Model 2			Model 3		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period	1.31	0.04	0.00	1.36	0.04	0.00	1.36	0.04	0.00
Unemployment two years ago	-0.48	0.06	0.00	-0.50	0.06	0.00	-0.50	0.06	0.00
Unemployment three years ago	0.04	0.03	0.29	0.08	0.03	0.02	0.07	0.03	0.03
Export market, 1st diff. (ΔXM_t)	-0.51	0.07	0.00	-0.46	0.07	0.00	-0.47	0.07	0.00
Export market, prev. period (XM_{t-1})	0.14	0.07	0.05	0.25	0.07	0.00	0.25	0.07	0.00
Export market, 1st diff. prev. period (ΔXM_{t-1})	-0.27	0.07	0.00	-0.29	0.08	0.00	-0.29	0.08	0.00
Change govt. purchases (dG_t)	-0.21	0.04	0.00	0.00	0.06	0.96	0.05	0.07	0.48
Interaction change govt. purchases (dG_t) and output gap $\bar{Y}_t$	0.07	0.01	0.00	0.14	0.02	0.00	0.19	0.05	0.00
Interaction change govt. purchases (dG_t) and output gap $\bar{Y}_t$, boom[a]				-0.18	0.04	0.00	-0.30	0.12	0.01
Interaction change govt. purchases (dG_t) and output gap $\bar{Y}_t^2$, boom[a]							0.02	0.02	0.40
Interaction change govt. purchases (dG_t) and output gap $\bar{Y}_t^2$, recession[a]							0.01	0.01	0.35
Obs = Country*Average groups	698	20	34.9	698	20	34.9	698	20	34.9
Standard deviation of residuals	0.63			0.65			0.65		
χ^2 of all the exogenous variables.[a]	272.76	(0.00)		255.85	(0.00)		257.60	(0.00)	
χ^2 of policy and exports.[a]	189.17	(0.00)		204.70	(0.00)		206.36	(0.00)	
1st order autocorrelation[a]	0.83	(0.40)		1.19	(0.23)		1.30	(0.20)	
2nd order autocorrelation[a]	-0.92	(0.36)		-1.16	(0.25)		-1.06	(0.29)	

Estimation method: Fixed effect coeffcients estimate, standard errors from GLS (xtgls without options).

Boom is a dummy for years where the output gap is postitive, and recession a dummy for negative output gap.

b) Numbers in parenthesis are p-values for the relevant null.

4.4.4 Government purchases, debt and openness

A key issue in part of the literature on the effect of fiscal policy is that the effect is likely to depend on private sector expectations on future fiscal policy. For example, Giavazzi and Pagano (1990) argue that a severe fiscal contraction might be expansionary in situations with concern for the risks of high public debt. This suggests that the effect of government purchases may depend on the level of public debt. In a recent study using structural VARs on quarterly data for 44 countries, both advanced and developing countries, Ilzetzki et al. (2010) find that the fiscal multiplier depends on the level of government debt, and that the fiscal multiplier is zero in high debt countries. To explore the possible importance of public debt, we interact the change in government purchases with lagged public debt as a ratio to GDP. To facilitate the interpretation, we measure the debt ratio as deviation from sample mean, which is equal to 0.6. We also include debt as a separate explanatory variable, as the levels of debt might well be correlated with the level of unemployment, cf. Bertola (2010).

The results are shown in model 2 in table 8. Surprisingly, the interaction term has the opposite sign of the expected, suggesting that the negative effect of a rise in government purchases on unemployment is larger in a country with high public debt. However, the result is only marginally significant, with a p-value of 7 percent. More importantly, as we shall see below, one may question whether the interaction term really captures the effect of public debt, or whether there are other possible explanations.

In table 8, we also explore whether the effect of government purchases depends on the openness of the country. According to traditional Keynesian analysis, the government expenditure multiplier is smaller in an open economy. In line with this, Beetsma and Giuliodori (2010) find in an analysis of 14 *EU* countries a clear positive effect of a rise in government purchases on GDP in "closed economies" (defined as countries where the ratio of export plus import to GDP is above sample average), and no significant effect in the remaining "open economies". Ilzetzki et al. (2010) also find stronger expansionary effect of open economies than in closed. To analyse the effect of openness, we interact the change in government purchases with an indicator of openness, based on the ratio of export plus import to GDP. As the degree of openness has increased over time, we consider two different specifications of this indicator. In model 3, the indicator measures the deviation of the export plus import ratio from the sample mean, implying that the indicator also captures the increase in openness over time. In model 4, the indicator is measured as deviation from year mean, thus omitting the change in openness over time. However, the effect is essentially the same in both specifications, indicating that a rise in government purchases has a stronger negative effect on unemployment in more open economies. This is the opposite effect of what one would expect from theory. However, as we shall see below, this effect does not hold up when we control for monetary regime.

Table 8: Equation (1) including openness in the economy

	Model 1			Model 2			Model 3			Model 4		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period	1.28	0.04	0.00	1.31	0.04	0.00	1.27	0.03	0.00	1.27	0.03	0.00
Unemployment two years ago	-0.39	0.05	0.00	-0.45	0.06	0.00	-0.39	0.05	0.00	-0.39	0.05	0.00
Unemployment three years ago	-0.01	0.03	0.82	0.01	0.04	0.71	-0.00	0.03	0.97	-0.00	0.03	0.98
Demand components:												
Export market, 1st diff. (ΔXM_t)	-0.53	0.06	0.00	-0.53	0.07	0.00	-0.54	0.06	0.00	-0.54	0.06	0.00
Export market, prev. period (XM_{t-1})	0.17	0.07	0.01	0.20	0.07	0.00	0.18	0.07	0.01	0.17	0.07	0.01
Export market, 1st diff. prev. period (ΔXM_{t-1})	-0.30	0.07	0.00	-0.30	0.08	0.00	-0.30	0.07	0.00	-0.30	0.07	0.00
Change govt. purchases (dG_t)	-0.19	0.04	0.00	-0.28	0.05	0.00	-0.19	0.04	0.00	-0.18	0.04	0.00
Interaction change govt. purchases and debt ratio prev. period (deviation from global mean), ($dG_t * (debt_{t-1} - \overline{debt})$)				-0.27	0.15	0.07						
Govt. debt prev. period ($debt_{it-1}$)				-0.19	0.19	0.30						
Interaction govt. purchases and openness deviation from sample mean prev. period ($dG_t * (open_{it-1} - \overline{open})$)							-0.20	0.09	0.03			
Openness prev. period ($open_{it-1}$)							-0.29	0.24	0.23	-0.34	0.23	0.14
Interaction govt. purchases and openness deviation from year mean prev. period ($dG_t * (open_{it-1} - \overline{open_t})$)										-0.25	0.10	0.02
Obs = Country*Average groups	801	20	40.0	643	20	32.1	801	20	40.0	801	20	40.0
Standard deviation of residuals	0.65			0.63			0.64			0.64		
χ^2 of all the exogenous variables.[a]	287.57	(0.00)		262.69	(0.00)		300.12	(0.00)		301.61	(0.00)	
χ^2 of policy and exports.[a]	195.35	(0.00)		186.59	(0.00)		206.83	(0.00)		208.20	(0.00)	
1st order autocorrelation[a]	0.65	(0.52)		1.41	(0.16)		0.95	(0.34)		0.95	(0.34)	
2nd order autocorrelation[a]	0.15	(0.88)		0.14	(0.89)		-0.03	(0.97)		-0.03	(0.97)	

Estimation method: Fixed effect coeffcients estimate, standard errors from GLS (xtgls without options).

a) Numbers in parenthesis are p-values for the relevant null.

4.4.5 Does the effect of government purchases vary over time?

In this subsection we investigate whether the effect of government purchases on unemployment varies over time, by allowing for a different effect for each decade. This exercise entails the added benefit that it facilitates comparison of our results with other studies on shorter sample periods. Model 2 in table 9 shows a striking difference. In the 1960s and 70s, we find essentially no effect of government purchases on unemployment, with coefficient estimates quite close to zero. In contrast, in the 1980s, 90s, and 2000s, the effect is much stronger than in the total sample, with coefficient estimates varying from -0.27 in the 2000s to -0.45 in the 1990s.

One might speculate that the absence of any effect in the 1960s reflects that unemployment in almost all countries was very stable and low, not giving much room for an effect of government purchases. In contrast, in the 1970s, unemployment rose quite sharply in most countries, and some countries tried to counteract this rise by use of expansionary fiscal policy. Thus, there could be a downward bias in the estimate reflecting that the rise in unemployment induced increased government spending. This story of reverse causality for the 1970s suggests the use of instrumental variables. However, the IV results in the third model in table 9 are rather consistent with the FE results; no effect of government purchases in the 1960s and 70s, and a negative effect for the last three decades, although with only a p-value of 15 percent in the 1980s. For the 2000s, the IV point estimate is almost three times as large as the FE estimate.

Kirchner et al. (2010) consider the effect of government purchase shocks in the aggregate euro area for the period 1980-2008. Using a time-varying structural VAR, they find that the short-run effectiveness of government purchases in stabilizing real GDP and private consumption has increased until the end of 1980s, but has decreased thereafter. Kirchner et al. (2010) also find that the multipliers in a longer horizons have declined substantially over their sample period.

In model 4, we include government debt in the FE regression. The motivation is to explore whether the unexpected sign for the debt-government purchases interaction in table 8 is related to the fact that the level of debt has varied over time, cf. table A1 in appendix 4.A. It turns out that when we allow the effect of government purchases to differ across decades, the debt-government purchases interaction loses its explanatory power, with a coefficient value close to zero with a p-value of 0.74. This suggests that the significant debt-government purchases interaction in table 8 is spurious, implying that we do not find any effect of the debt level on the impact of government purchases on unemployment.

Table 9: Equation (1) for each decade in the sample

	Model 1			Model 2			IV 3[a]			Model 4		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period	1.28	0.04	0.00	1.24	0.03	0.00	1.28	0.05	0.00	1.27	0.04	0.00
Unemployment two years ago	-0.39	0.05	0.00	-0.38	0.05	0.00	-0.46	0.07	0.00	-0.44	0.06	0.00
Unemployment three years ago	-0.01	0.03	0.82	-0.02	0.03	0.57	-0.01	0.04	0.75	0.00	0.04	0.94
Demand components:												
Export market, 1st diff. (ΔXM_t)	-0.53	0.06	0.00	-0.48	0.06	0.00	-0.46	0.07	0.00	-0.50	0.07	0.00
Export market, prev. period (XM_{t-1})	0.17	0.07	0.01	0.16	0.07	0.03	0.17	0.08	0.04	0.16	0.07	0.03
Export market, 1st diff. prev. period (ΔXM_{t-1})	-0.30	0.07	0.00	-0.28	0.07	0.00	-0.17	0.09	0.05	-0.24	0.08	0.00
Change govt. purchases (dG_t)	-0.19	0.04	0.00									
Change govt. purchases (dG_t), 1960s				0.04	0.11	0.69	0.12	0.27	0.67	0.04	0.18	0.83
Change govt. purchases (dG_t), 1970s				0.04	0.06	0.50	0.28	0.38	0.46	0.04	0.09	0.63
Change govt. purchases (dG_t), 1980s				-0.33	0.08	0.00	-0.42	0.29	0.15	-0.48	0.10	0.00
Change govt. purchases (dG_t), 1990s				-0.45	0.09	0.00	-0.38	0.22	0.08	-0.44	0.09	0.00
Change govt. purchases (dG_t), 2000s				-0.28	0.09	0.00	-0.81	0.29	0.01	-0.32	0.09	0.00
Dummy for 1970s				0.25	0.18	0.17	0.23	0.46	0.62	0.32	0.26	0.23
Dummy for 1980s				0.77	0.19	0.00	1.13	0.38	0.00	0.97	0.27	0.00
Dummy for 1990s				0.89	0.20	0.00	1.22	0.37	0.00	1.09	0.27	0.00
Dummy for 2000s				0.65	0.20	0.00	1.25	0.40	0.00	0.84	0.28	0.00
Interaction change govt. purchases and debt ratio												
(deviation from global mean) prev. period, ($dG_t * (debt_{t-1} - \overline{debt})$)										-0.06	0.16	0.74
Lagged levels of govt. debt										-0.31	0.20	0.12
Obs = Country*Average groups	801	20	40.0	801	20	40.0	626	20	31.3	643	20	32.1
Standard deviation of residuals	0.65			0.63			0.98			0.61		
χ^2 of all the exogenous variables.[b]	287.57	(0.00)		275.60	(0.00)		186.28	(0.00)		263.00	(0.00)	
χ^2 of policy and exports.[b]	195.35	(0.00)		198.66	(0.00)		106.02	(0.00)		184.38	(0.00)	
1st order autocorrelation[b]	0.65	(0.52)		0.98	(0.33)					1.26	(0.21)	
2nd order autocorrelation[b]	0.15	(0.88)		0.26	(0.79)					0.56	(0.58)	

Estimation method: Fixed effect coeffcients estimate, standard errors from GLS (xtgls without options) is used in all the regressions except for in model 3 which IV approach is used.

In all equations it is also controlled for labour market institutions.

a) Change govt. purchases (dG_t) is treated as endogenous. Instruments are: (ΔdG_{t-1}), (dG_{t-2}) and ($\Delta debt_{t-1}$).

b) Numbers in parenthesis are p-values for the relevant null.

4.4.6 Distinguishing between types of government purchases: investment, wage consumption and non-wage consumption

Both from a theoretical and policy perspective it is of considerable interest to explore whether the effect of a change in government purchases differs depending of the type of purchase. In statistical sources, one typically distinguishes between three categories, which we also use in our analysis: government wage consumption, which is essentially public employment ($dCGW$), government non-wage consumption ($dCGNW$) and government real investments (dIG). We consider the same form of the left hand side variable as before, i.e. the change in each of this categories (indicated by the d in the variable name), in real terms, and measured as share of trend-GDP, see appendix 4.A for a detailed explanation.

The results are presented in table 10. Model 1 shows that government investments and government wage consumption both have a significant negative impact on the unemployment rate (although government invest only with a p-value of 0.08), while the estimated effect of government non-wage consumption is close to zero and not statistically significant.

In model 2 in table 10, we explore whether the effect of the different types of government purchase depends on the cyclical state of the economy. The interaction terms for both government investment and government wage consumption are positive and statistically significant, implying that both increased government investment and increased government wage consumption have a stronger dampening effect on unemployment when the output gap is negative, consistent with our prior results.

The third model in table 10 presents the result of model 1 using IV; we find that the effect of government investment is much stronger, with a point estimate of -0.59, and a p-value of 0.05, while the other coefficients are not statistically significant.

Ilzetzki et al. (2010) also explore possible differences between the effect of government consumption and government investment, and they find about the same point estimates for both consumption and investment for advanced countries, with multiplier estimates (the effect of government purchases on GDP) of 0.4 at impact and 0.8 in the long run.

Table 10: Equation (1) with different types of government purchases and interaction with the output gap

	Model 1			Model 2			IV 3[a]		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period	1.31	0.04	0.00	1.30	0.04	0.00	1.25	0.06	0.00
Unemployment two years ago	-0.39	0.06	0.00	-0.43	0.07	0.00	-0.37	0.07	0.00
Unemployment three years ago	-0.01	0.04	0.83	0.03	0.04	0.41	-0.04	0.04	0.32
Demand components:									
Export market, 1st diff. (ΔXM_t)	-0.52	0.07	0.00	-0.48	0.07	0.00	-0.51	0.08	0.00
Export market, prev. period (XM_{t-1})	0.26	0.08	0.00	0.24	0.08	0.00	0.19	0.09	0.04
Export market, 1st diff. prev. period (ΔXM_{t-1})	-0.33	0.08	0.00	-0.32	0.09	0.00	-0.29	0.10	0.00
Change govt. investments, (dIG_t)	-0.11	0.06	0.08	-0.13	0.07	0.06	-0.59	0.30	0.05
Change govt. non-wage consumption, ($dCGNW_t$)	0.01	0.01	0.73	0.00	0.02	0.78	-0.01	0.07	0.87
Change govt. wage consumption, ($dCGW_t$)	-0.29	0.09	0.00	-0.27	0.09	0.00	-0.25	0.33	0.45
Interaction dGI_t and $\bar{Y}_t$				0.08	0.03	0.02			
Interaction $dCGNW_t$ and $\bar{Y}_t$				0.01	0.01	0.37			
Interaction $dCGW_t$ and $\bar{Y}_t$				0.14	0.03	0.00			
Obs = Country*Average groups	595	16	37.2	537	16	33.6	501	16	31.3
Standard deviation of residuals	0.66			0.64			0.84		
χ^2 of all the exogenous variables.[b]	237.52	(0.00)		244.56	(0.00)		185.35	(0.00)	
χ^2 of policy and exports.[b]	167.66	(0.00)		181.62	(0.00)		101.43	(0.00)	
1st order autocorrelation[b]	1.30	(0.19)		1.39	(0.16)				
2nd order autocorrelation[b]	0.47	(0.64)		0.12	(0.91)				

Estimation method: Fixed effect coeffcients estimate, standard errors from GLS (xtgls without options)
is used in all the regressions except for in model 3 which IV approach is used.
In all equations it is also controlled for labour market institutions.
a) Change govt. purchases (dG_t) is treated as endogenous. Instruments are: (ΔdG_{t-1}), (dG_{t-2}) and ($\Delta debt_{t-1}$).
b) Numbers in parenthesis are p-values for the relevant null.

Table 11: Decomposition of government purchases, as share of total government purchases (sample average)

	Mean	Std.	Obs.
Government wage consumption CGW	0.54	0.09	772
Government non-wage consumption CGNW	0.29	0.08	772
Government investments IG	0.17	0.08	772

4.4.7 Government purchases and monetary regime

In this subsection we explore whether the effect of government purchases depends on the monetary regime, as implied by standard text book macro like the Mundell Fleming model, and also emphasized in much of the recent literature, e.g. Ilzetzki et al. (2010). Under an inflation target, an expansionary effect of increased government purchases will be counteracted by a rise in the interest rate, partly offsetting the effect on unemployment. Also with other types of floating exchange rates, one would expect an expansionary effect from fiscal policy be counteracted by changes in the exchange rate and the interest rate. In contrast, if the nominal interest rate is unaffected, as it will be with a fixed exchange rate and for a small country in a monetary union, and inflation increases so that the real interest falls, the government multiplier might be considerably above unity (see e.g. Christiano et al. (2009).

We use four dummies to capture the different monetary regimes within the sample period; the Bretton Woods agreement (all countries until 1972), a fixed exchange rate regime, a floating exchange rate regime, and membership in the European Monetary Union (EMU). In the Bretton Woods agreement, all currencies were tied to US dollars. Floating exchange rate includes various forms of floating, and in recent years also inflation targeting. Countries that took part in the European Exchange Rate Mechanism ERM are defined as having a fixed exchange rate regime, except for Germany, which we define as floating in light of Germany's dominating position and the independent status of the Bundesbank. We also tried to distinguish between credible and non-credible fixed exchange rate regimes depending on the interest rate differential relative to the anchor country (in most cases Germany), defining the regime as non-credible if the interest rate differential exceed 1 percentage point in annual terms. The idea here is that if the fixed exchange rate lacks credibility, a fiscal expansion could have a negative effect on the economy by impairing credibility, for example raising devaluation expectations and thus also push up interest rates. However, the point estimates were essentially the same for credible and non-credible fixed exchange rate regimes, so we decided to drop this distinction in the results we report. Further description of the monetary regimes is given in appendix 4.A.

Table 12: Estimation of equation (1) with different types of monetary regime

	Model 1			Model 2			Model 3			Model 4		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period	1.24	0.03	0.00	1.24	0.03	0.00	1.23	0.03	0.00	1.23	0.03	0.00
Unemployment two years ago	-0.37	0.05	0.00	-0.37	0.05	0.00	-0.37	0.05	0.00	-0.37	0.05	0.00
Unemployment three years ago	-0.01	0.03	0.77	-0.01	0.03	0.77	-0.00	0.03	0.96	-0.00	0.03	0.95
Demand components:												
Export market, 1st diff. (ΔXM_t)	-0.51	0.06	0.00	-0.51	0.06	0.00	-0.53	0.06	0.00	-0.53	0.06	0.00
Export market, prev. period (XM_{t-1})	0.15	0.07	0.03	0.15	0.07	0.03	0.10	0.07	0.13	0.10	0.07	0.14
Export market, 1st diff. prev. period (ΔXM_{t-1})	-0.29	0.07	0.00	-0.29	0.07	0.00	-0.26	0.07	0.00	-0.26	0.07	0.00
Change govt. purchases (dG_t), Bretton woods	0.01	0.08	0.86	0.04	0.10	0.67	0.03	0.08	0.76	0.02	0.08	0.79
Change govt. purchases (dG_t) Monetary union (EMU)	-0.35	0.11	0.00	-0.32	0.13	0.01	-0.27	0.16	0.08	-0.29	0.14	0.04
Change govt. purchases (dG_t), Fixed exchange rate	-0.45	0.07	0.00	-0.42	0.09	0.00	-0.46	0.07	0.00	-0.47	0.07	0.00
Change govt. purchases (dG_t), Floating exchange rate	-0.03	0.06	0.64				-0.04	0.06	0.54	-0.03	0.06	0.64
Dummy for Bretton woods	-0.26	0.14	0.06	-0.26	0.14	0.06	-0.40	0.15	0.01	-0.40	0.15	0.01
Dummy for EMU	0.14	0.15	0.33	0.14	0.15	0.33	0.36	0.17	0.04	0.38	0.17	0.02
Dummy for Fixed exchange rate	0.47	0.12	0.00	0.47	0.12	0.00	0.54	0.12	0.00	0.55	0.12	0.00
Change govt. purchases, (dG_t)				-0.03	0.06	0.64						
Interaction change govt. purchases and openness deviation from sample mean prev. period ($dG_t * (open_{t-1} - \overline{open})$)							0.04	0.13	0.74			
Openness prev. period ($open_{it-1}$)							-1.08	0.28	0.00	-1.09	0.28	0.00
Interaction change govt. purchases and openness, deviation from year mean prev. period ($dG_t * (open_{t-1} - \overline{open_t})$)										0.09	0.13	0.49
Obs = Country*Average groups	801	20	40.0	801	20	40.0	801	20	40.0	801	20	40.0
Standard deviation of residuals	0.63			0.63			0.63			0.63		
χ^2 of all the exogenous variables.[a]	305.72	(0.00)		305.72	(0.00)		304.26	(0.00)		304.78	(0.00)	
χ^2 of policy and exports.[a]	218.72	(0.00)		218.72	(0.00)		210.29	(0.00)		210.76	(0.00)	
1st order autocorrelation[a]	-0.29	(0.77)		0.75	(0.45)		0.51	(0.61)		0.34	(0.74)	
2nd order autocorrelation[a]	0.21	(0.83)		0.25	(0.80)		0.19	(0.85)		0.19	(0.85)	

In all equations it is also controlled for labour market institutions.

a) Numbers in parenthesis are p-values for the relevant null.

Model 1 in table 12 shows that the effect of government purchases differs sharply across monetary regimes. The point estimate is −0.35 in the EMU and −0.45 with a fixed exchange rate regime, and highly statistically significant. In contrast, during the Bretton Woods regimes, and with a floating exchange rate, the point estimate is essentially zero. These results are in line with our theoretical expectations, where fiscal policy is effective under a fixed exchange rate regime, but not under float. The exception is of course the Bretton Woods period; however, this finding only reflects the prior finding of no effect of fiscal policy during the 1960s and 70s. The difference across exchange rate regimes is consistent with those of Ilzetzki et al. (2010); they find a significant positive effect of increased government consumption on GDP for fixed exchange rate regimes, and while the effect is significant and negative at impact for floating regimes.

Model 2 shows that the difference between regimes is indeed statistically significant. In model 3 and 4, we review our prior findings on the link between fiscal policy and the openness of the economy. We find that the interaction between openness and the change in government purchases is no longer significant. Indeed, for both specifications the point estimate is close to zero, suggesting that our prior findings on openness are spurious and caused by a correlation between openness and monetary regime.

4.5 Concluding remarks

The vast changes in budget policies in most OECD countries in recent years have made it compelling for economists to find out what the effects will be on the economy. Politicians would like to know the consequences on the GDP, unemployment and public debt, etc, of a reduction in public expenditure, and economists have struggled with providing an answer. In principle, there should be a clear objective answer to questions like this, even if the answer obviously will depend on the specific situation, like type of public expenditure, monetary policy response, the effect on private sector expectations, etc. However, there is no agreement in the literature about what this answer should be.

After decades of little research on the effects of fiscal policy, there is now a lot of recent studies exploring this issue. Most of the studies are based on structural VARs, or analyze the effects within structural macro models. As these approaches have their strengths and weaknesses, it is of value also to try other methods. We investigate the effect of changes in government purchases on unemployment by use of panel data estimation, building on an empirical equation where long run unemployment is a function of a number of labour market variables, along the lines of Layard et al. (2005) and Nickell et al. (2005). One advantage with this approach is that we are able to control for key labour market variables, which according to a large literature has been important for the evolution of OECD unemployment.

We find that an increase in government purchases has an economically and statistically significant dampening effect on unemployment. According to our base specification, a permanent increase in government purchases equal to one percent of GDP on average leads to a reduction in unemployment of 0.2 percentage point, increasing to 0.25 percentage points after one year, for then to gradually vanish over the following decade. Instrumental variable estimation suggests that this estimate is downward biased, and the IV point estimate is a reduction in unemployment of 0.5 percentage point following an increase in government purchases equal to one percent of GDP. There is considerable variation in the effect of government purchases, across countries and also across time periods and depending on other specific circumstances. We find no effect in the 1960s and 70s, and a correspondingly stronger than average effect in the 1980s, 90s and 2000s. We also find that the effect is considerably larger in a weak cyclical situation; when the output gap is equal to minus three percent, the effect on unemployment is about double of the average effect. The monetary regime is important for the effect. In line with the Mundell Fleming model, we find a strong effect of government purchases on unemployment for countries within a monetary union or with a fixed exchange rate regime (excluding the Bretton-Woods), and no significant effect of government purchases for countries with a floating exchange rate. Distinguishing between different types of government purchases, we only find a strong significant effect of government wage consumption (i.e. public employment), and to some extent also government investment (with p-value of 5 percent), but not of government non-wage consumption.

4.A The Data: Definitions and sources

The data used in this paper are explained in this appendix. The sample period is from 1960 to 2007. The countries in the sample are:

Australia	Finland	Japan	Spain
Austria	France	Netherlands	Sweden
Belgium	Germany	Norway	Switzerland
Canada	Ireland	New Zealand	United Kingdom
Denmark	Italy	Portugal	United States

4.A.1 Government purchases

The change in government purchases (dG) is measured as the growth rate in real terms of government purchases, multiplied with government purchases as a share of trend GDP. The data are from OECD (2008b) unless otherwise noted. The formula of dG is:

$$dG_{it} = \frac{(CGV_{it} + IGV_{it}) - (CGV_{it-1} - IGV_{it-1})}{(CGV_{it-1} + IGV_{it})} * \frac{CG_{it} + IG_{it} - CFKG_{it}}{YCT_{it}} * 100 \quad \text{(A1)}$$

where CG is government consumption, IG government investments, $CFKG$ is consumption of fixed capital, and YCT is trend-GDP. The variables are in nominal prices, except those where the last letter V indicates real terms. Note that government purchases do not include transfers like social security expenditures etc. Note also that we subtract consumption of fixed capital ($CFKG$) from government consumption to obtain the actual expenditure, as the consumption of fiscal capital is an imputed measure. $CFKG$ is not subtracted in the real growth rate for reasons of data availability, but this is unimportant as there presumably is little variation over time in the imputed consumption of fixed capital. Investment data is missing for some countries (Spain, Italy, Switzerland) and for these countries we use government consumption.

Trend-GDP is equal to the backward looking 10 year moving average of real GDP (YQ) multiplied with the two year moving average of the price deflator ($PGDP$) to a variable in nominal terms. Real GDP is prolonged backwards (1950-1960) with the growth rate in GDP in The Conference Board (2010). We have used the GDPGK series expressed in 1990 U.S. dollars. Germany is prolonged backwards with the sum of West Germany and East Germany before 1989, and the data for East Germany is linearly interpolated when observations are missing.

The variable IGV is calculated as $IGV = IG/PIG$, where PIG is the associated price deflator.

The change for the categories of government purchases; $dCGW$ government purchases for wage-consumption, $dCGNW$ non-wage-consumption and dIG investments in fixed capital, are calculated by use of the formula above. The variables in real terms are when necessary (i.e. because they are not supplied by the OECD) calculated as above; real government wage consumption ($CGW/PCGW$), real government non-wage consumption ($CGNW/PCGNW$) and real gross government investments (IG/PIG), where $PCGW$, $PCGNW$, *and* PIG are the associated price deflators. There was no data available for $PCGNW$, which was then calculated from the identity $CG = CGW + CGNW$, which should also hold in real terms, $CG/PCG = CGW/PCGW + CGNW/PCGNW$, or $PCGNW = CGNW/(CG/PCG - CGW/PCGW)$. Some clearly implausible values for the growth rate of $CGNW$ were dropped (287 for Spain in 1983, 140 and 34 in United Kingdom in 1967 and 1970, a fall in the same variable in United Kingdom in 1968 and 1969 of 24 and 27, and finally a fall equal to 20 in Ireland in 1971). Investment data is missing for some countries (Spain, Italy, Switzerland) and for these countries the missing observations are set to zero.

4.A.2 Output gap

The output gap is in percentage points and collected from OECD (2008b). The output gap and the government purchases indicator are illustrated in figure 1 and 2.

4.A.3 Monetary regime

We have constructed 5 dummies to account for changes in the monetary regime over the sample period; the Bretton Woods agreement (until 1972), a floating exchange rate, a credible fixed exchange rate, a non-credible fixed exchange rate, and membership in the European Monetary Union (EMU). The dummy $D_{bretton} = 1$ indicates the Bretton Woods agreement covering all countries in the sample in the period 1960 to 1972. In the Bretton Woods agreement, all currencies were tied to US dollars. The dummy D_{float} indicates that a number of countries adopted a floating exchange rate from 1973: Australia, Canada, Germany, Japan, New Zealand, United States and United Kingdom (except 1990 and 1991 [3]), later also Sweden (since 1992) and Norway (since 1999) adopted a floating exchange rate. This regime includes various forms of floating. In recent years also inflation targeting. The dummy D_{cred} indicates a credible fixed exchange rate, and $D_{noncred}$ a the non-credible one. Countries that took part in the European Exchange Rate Mechanism ERM are defined as having a fixed exchange rate regime, except for Germany, which we define as floating in light of Germany's dominating position and the independent status of the Bundesbank. The distinction between credible and non-credible fixed exchange rate regimes is based on the interest rate differential relative to the anchor country (in most cases Germany), where the regime is defined as non-credible if the interest rate differential

[3] The UK was a member of the European exchange rate mechanism (ERM) from October 1990 to September 1992

exceed 1 percentage point in annual terms, or if the regime is non-credible in subsequent years before entry into the monetary union. In other words, a fixed exchange rate regime is only defined as credible if the interest rate differential is less than one percent in all the following years. D_{EMU} indicates EMU membership, covering Austria, Belgium, Finland, France, Germany, Ireland, Italy, Portugal and Spain since 1999. Figure A1 and A2 illustrate the dummies for monetary regime for each of the countries in the sample.

4.A.4 Export market indicator

The export market (XM) indicator is calculated as a weighted average of the GDP-gap of the trading partners, where the GDP-gap is the deviation of GDP from Hodrick Prescott-trend (with smoothing parameter 100), and the weights reflect the share of the exports from country i that goes to each of the trading partners j. The export market indicator is displayed in figure A3 and A4. The formulae is

$$XM_{it} = \Sigma_j w_{ijt} * GAP_{jt} \tag{A2}$$

where $w_{ijt} = x_{ijt}/\Sigma_j x_{ijt}$. x_{ijt} is export from country i to county j in year t. GAP_{jt} is the GDP-gap in country j in year t. The trading partners to one country in the sample are all the other countries in the sample and the rest of 'the world'. 'The world' is one country's total export subtracted the sum of exports to all countries in the sample. The The exports data is from SITC Revision 2 OECD (2010), and are used to calculate the export shares for each country in the sample.

The exports from Germany includes Eastern Germany since 1991 and as a partner country Eastern Germany is included the whole sample period. The time series are prolonged backwards with the exports to the world when observations are missing. For instance if Australia has exported to Canada, but exports are only reported in the years 2003 and onwards, the exports from Australia to Canada is prolonged backwards with the change in Australia's world export growth rate and is equal to the total export for Australia. The same method is used for extracting the time series forward. For Belgium the exports data is only reported from 1993. Before 1993 the data is constructed by prolonging the export data for Belgium by use of the change in exports for the Belgium-Luxembourg Economic Union.

The GDP-gap for the XM indicator is calculated as the deviation of GDP from trend GDP, divided by trend GDP. Trend GDP derived by use of Hodrick and Prescott (1997) filter, (the HP filter hereafter). The value of the smoothing parameter has been discussed in several papers in range of 6.25 up to 400, see Backus and Kehoe (1992), Correia et al. (1992) and Baxter and King (1999). We would like to remove the difference in growth rates from 1960 to 2007, but not remove cycles, irrespective of whether they are caused by business cycle movements or structural changes in the economy. Therefore we use a

Figure A1: Monetary regime, interest rate differential and the output gap during the period 1960 to 2007.

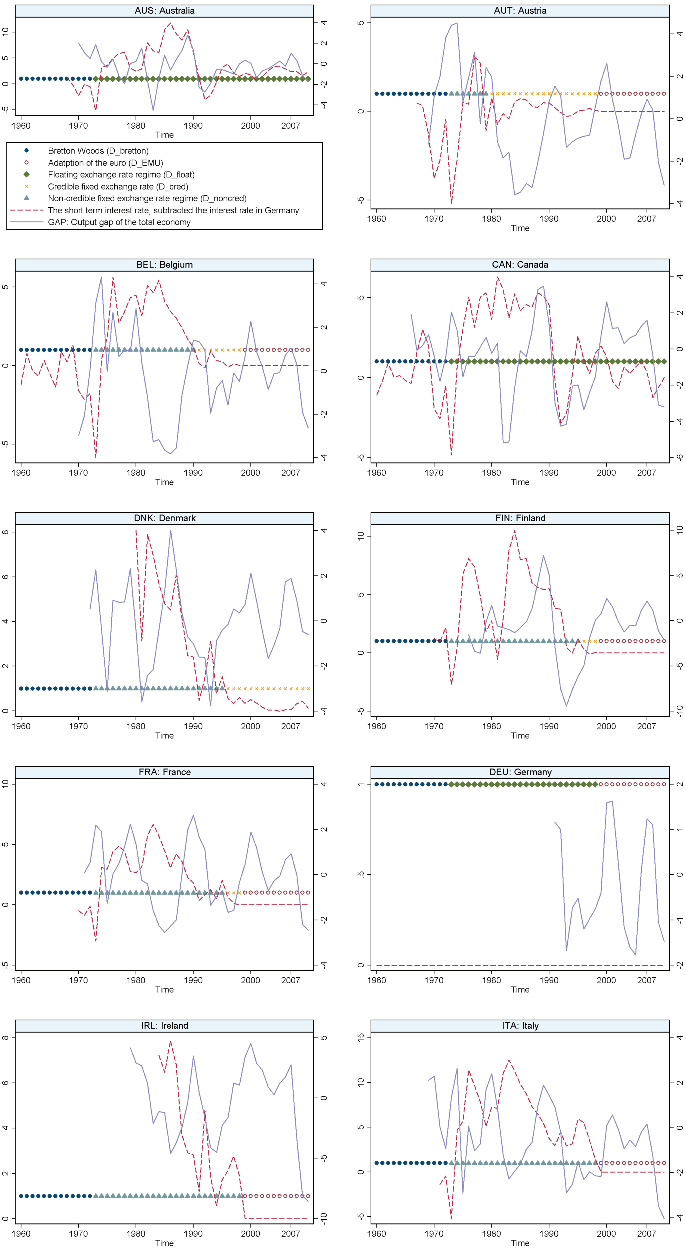
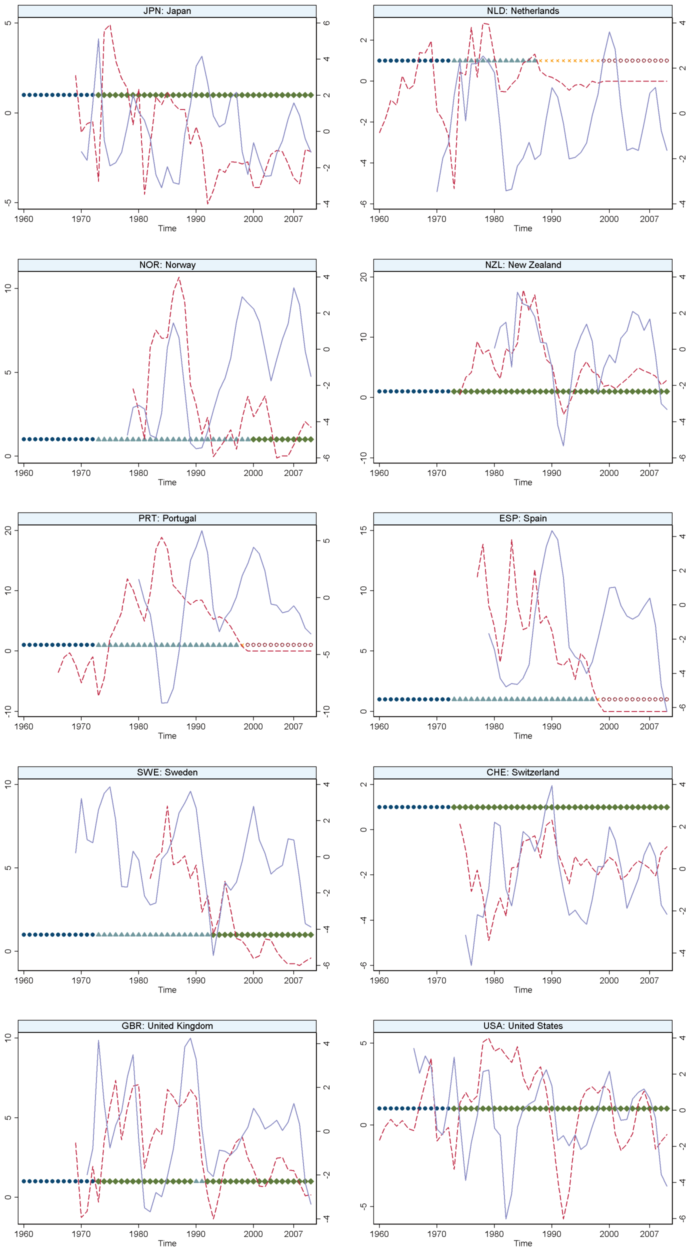


Figure A2: Monetary regime, interest rate differential and the output gap during the period 1960 to 2007.



rather high value of the smoothing parameter equal to 100, this is also in line with Backus and Kehoe (1992).

The GDP-gap for each of the twenty OECD countries is calculated using data from OECD (2008b). The world GDP-gap is constructed using data for the real GDP in The Conference Board (2010). We have used the $GDPGK$-series with are GDP expressed in 1990 U.S. dollars, and covers 123 countries in the database. The world GDP is the sum of the 123 countries subtracted the value of GDP in each of the twenty countries in the sample. Note, that East Germany is linearly interpolated when observations are missing, and Germany is prolonged backwards with the sum of West Germany and East Germany before 1989. The aggregate of former Union of Soviet Socialist Republics, Yugoslav federation and Czechoslovakia is used in the world GDP due to missing observations for each of the countries separately.

4.A.5 Election year

The election year variable 'elect' is collected from Armingeon et al. (2010), and the original data source is European Journal of Political Research (Political Data Yearbook, various issues); Mackie and Rose (1991); Keesing's Archive; Parline database. The variable describes date of election of national parliament (lower house). The variable covers the years in the period 1960 to 2008.

4.A.6 Debt

The gross government debt variable, as a percent of GDP, 'debt', is collected from Armingeon et al. (2010), and the original data source is several versions of Oecd Economic outlook. See details regarding versions and the mission observations in Codebook by Armingeon et al. (2010). To facilitate the interpretation of the coefficients, the variable is rescaled to be the share of GDP. Gross government debt as a percentage of GDP is shown in table A1.

4.A.7 Openness

The openness variable is total trade (export and imports) in percentage of GDP. The variable is collected from Armingeon et al. (2010). See details regarding versions and the mission observations in Codebook by Armingeon et al. (2010). To facilitate the interpretation of the coefficients, the variable is rescaled to be the share of GDP. Gross government debt as a percentage of GDP is shown in table A1.

Figure A3: Exports market indicator and GDP-gap. Percentage points.

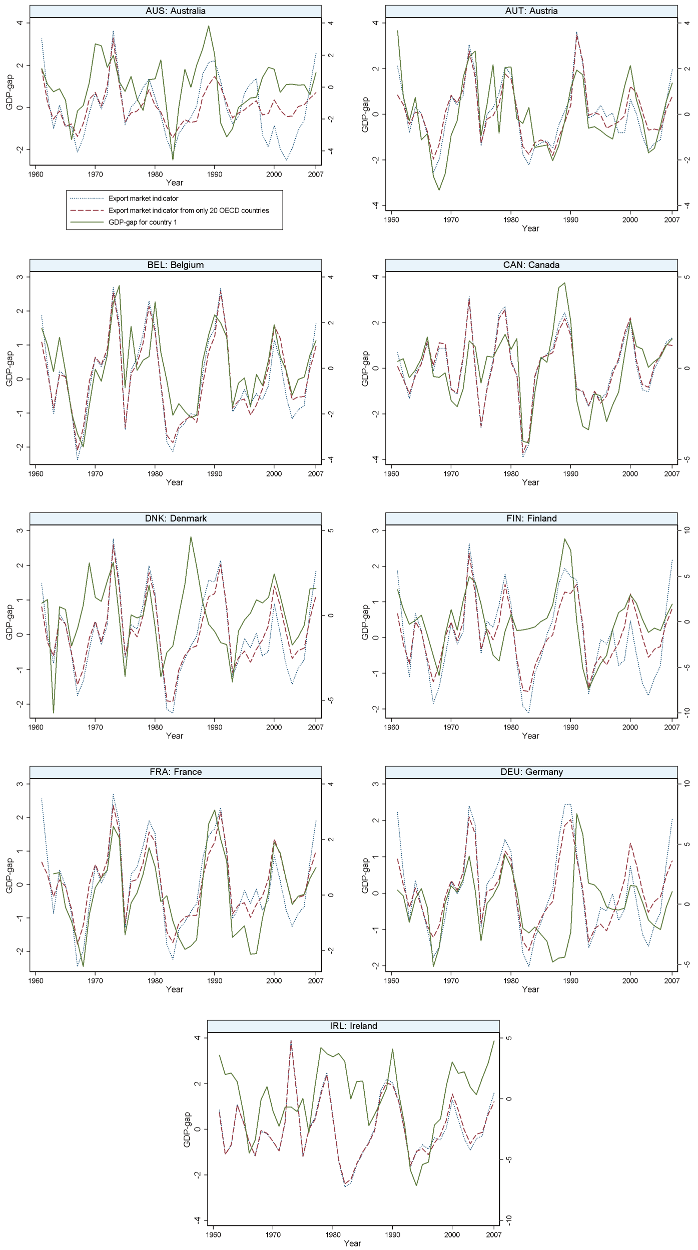


Figure A4: Exports market indicator and GDP-gap. Percentage points

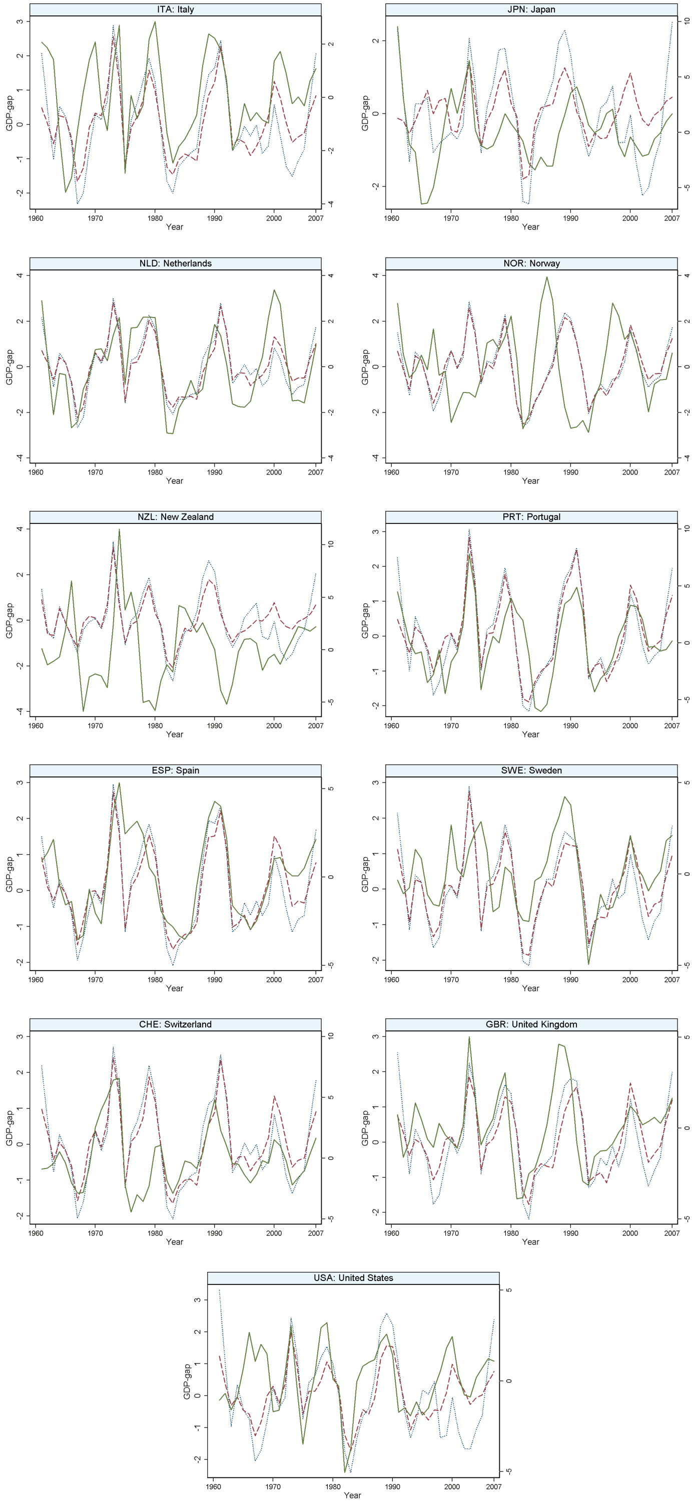


Table A1: Public debt for the countries in the panel over the sample period.

Country stats	**1960-69**	**1970-79**	**1980-89**	**1990-99**	**2000-07**	**1960-07**
Australia mean	.	.	0.24	0.33	0.19	0.26
sd	.	.	0.01	0.07	0.03	0.09
Austria mean	.	0.23	0.48	0.65	0.70	0.50
sd	.	0.07	0.09	0.06	0.04	0.19
Belgium mean	.	0.60	1.08	1.31	1.01	1.00
sd	.	0.04	0.17	0.07	0.09	0.28
Canada mean	0.61	0.48	0.62	0.93	0.75	0.68
sd	0.06	0.04	0.10	0.08	0.06	0.17
Denmark mean	.	.	0.66	0.73	0.48	0.63
sd	.	.	0.10	0.07	0.09	0.14
Finland mean	.	0.11	0.17	0.52	0.49	0.34
sd	.	0.03	0.02	0.18	0.04	0.20
France mean	.	0.33	0.35	0.57	0.70	0.48
sd	.	0.04	0.04	0.12	0.04	0.17
Germany mean	.	0.22	0.38	0.51	0.65	0.43
sd	.	0.05	0.04	0.10	0.04	0.17
Ireland mean	.	0.61	0.94	0.80	0.34	0.70
sd	.	0.05	0.15	0.16	0.04	0.26
Italy mean	0.37	0.75	0.91	1.18	1.18	0.91
sd	0.03	0.13	0.06	0.13	0.03	0.29
Japan mean	.	0.24	0.65	0.87	1.59	0.80
sd	.	0.12	0.10	0.22	0.14	0.50
Netherlands mean	0.72	0.56	0.79	0.86	0.59	0.71
sd	0.03	0.05	0.11	0.07	0.04	0.13
New Zealand mean	.	.	.	0.49	0.31	0.39
sd	.	.	.	0.09	0.04	0.11
Norway mean	.	0.43	0.34	0.34	0.47	0.39
sd	.	0.05	0.04	0.05	0.10	0.08
Portugal mean	.	.	.	0.66	0.69	0.68
sd	.	.	.	0.03	0.04	0.04
Spain mean	.	.	0.47	0.64	0.55	0.58
sd	.	.	0.02	0.11	0.08	0.11
Sweden mean	.	0.30	0.62	0.74	0.59	0.56
sd	.	0.03	0.09	0.13	0.06	0.19
Switzerland mean	.	.	.	0.45	0.54	0.49
sd	.	.	.	0.08	0.04	0.08
United Kingdom mean	0.95	0.58	0.47	0.45	0.44	0.53
sd	0.05	0.08	0.05	0.08	0.03	0.16
United States mean	0.56	0.44	0.52	0.68	0.59	0.56
sd	0.05	0.02	0.08	0.04	0.03	0.09
Total mean	0.62	0.43	0.59	0.69	0.64	0.60
sd	0.16	0.19	0.26	0.27	0.31	0.28

4.A.8 Unemployment rate

The unemployment variable is available at a yearly frequency in OECD (2008a)[4] and missing observations are replaced with observations from earlier data bases OECD (2002), OECD (2006) and OECD (2008b).[5]

The standardized unemployment rate (UNR) in Economic Outlook OECD (2008a) is used as a primary data source for the unemployment rate in the OECD countries, and missing observations are replaced by the growth rate in a corresponding time series in an earlier data base OECD (2002). Australia, Denmark, Germany, Spain and Switzerland are prolonged by the formula in equation (A3):

$$Y_{it} = Y_{it+1} * \frac{X_{it}}{X_{it+1}} \tag{A3}$$

where Y_it denotes (UNR) in OECD (2008a) and X_{it} denotes the (UNR) in the earlier data base OECD (2002) for country i in time period t.

Australia and Denmark are prolonged five years backwards. Germany from 1991, Spain from 1976, Switzerland from 1969 and backwards.

4.A.9 Labour market institutions

New information for institutional variables are available every second or fifth year. Labour market institutions such as the tax wedge, the union density, the coordination among wage setters, and the benefit replacement ratio and duration are used in this paper. The variables and the method of combining data sources are discussed in detail in the next sections.

TW: Tax wedge

The rates described here are calculated from actual tax payments. The total tax wedge is equal to the sum of the employment tax rate (t1), the direct tax rate (t2) and the indirect tax rate (t3), as given in Equation (A4).

$$TW = t1 + t2 + t3 \tag{A4}$$

$t1$ is equal to employers‘ total wage costs calculated by the sum of wages received by employees and taxes payed by the employer to the government. This gives the following relationship; $t1 = SSRG/(IE - SSRG)$, where SSRG is social security contributions and IE is compensation to employees. The latter two consist of two main components, wages and salaries and social contributions. Social contributions are payed by the employers to social security schemes or private funded social insurance schemes. $t2$, are direct taxes

[4]Data are collected and organized by the author. This implies that neither OECD nor any other source is responsible for the analysis or the interpretation of the data in this paper.

[5]An comprehensive overview of data and data sources is available upon request.

payed by the households (TAXh) divided by current receipts of households (CRh), i.e. $t2 = TAXh/CRh$. Finally $t3 = (TAXind - SUB)/Cp$, where TAXind are net indirect taxes, SUB is the value of subsidies and Cp is the value of private final consumption expenditure.

The main data source for tax wedges is OECD (2008c) which contains information for the period 1960 to 2010. The latter years are predictions. The tax rates are calculated by the above formulas, and when a tax rate is missing, the growth rate in the same tax rate but from the data base of Nickell (2006) in the period 1960 to 2003 is used to prolong the time series for the following countries: Belgium is prolonged before 1965, Denmark is prolonged before 1966, Germany before 1970, Portugal is prolonged in the period 1960 to 1995 and Switzerland is prolonged before 1990 with the tax rates in OECD (2008c). Tax rates for Australia, Austria, Canada, Finland, France, Ireland, Italy, Japan, Netherlands, Norway, Spain, Sweden, the United Kingdom and the United States are not prolonged and taken directly from the main data source OECD (2008c). New Zealand has main data source Nickell et al. (2005) for the period 1975 to 1986 due to missing observations in OECD (2008c). Time series for $t1$, $t2$ and $t3$ from Nickell (2006) is used to extend the main data source: The growth rate of the sum of $t1$ and $t2$ is used before 1975, and the growth rate in $t3$ after 1986. Note also that the $t3$ is interpolated due to one missing observation in 1991.

BRR: Benefit replacement rates

Benefit replacement ratio is a measure of how much each unemployed worker receives in benefits from the government. The benefit replacement ratio is described in detail below.

The detailed rate for unemployment benefits divides data in three different family types: single, with a dependent spouse and with a working spouse. The benefits also depend on the employment situation: 67 percent and 100 percent of the average earnings. Within these groups, benefits are divided into the duration of benefits when being unemployed. One variable for how much each of the former groups receives in the first year, the second and third year and the fourth and fifth year. The indexes used in this paper uses the indexes aggregated over over family types. This results in six different groups: brr67a1, brr67a2, brr67a4, brr100a1, brr100a2 and brr100a4.

brr67a1: First year benefit replacement rate for workers with 67 percent of average earnings and the average over family types.

brr67a2: Benefit replacement for the second and third year. 67 percent of average earnings and the average over family types.

brr67a4: Benefit replacement rate for the fourth and fifth year. 67 percent of average earnings and the average over family types.

brr100a1, brr100a2 and brr100a4: The same as the former but for 100 percent of average earnings.

The average of brr67a1 and brr100a1 is used as an indicator of benefit replacement ratios, i.e. BRR.

The main source for the more detailed benefit ratios are tables in employment outlook, see OECD (2004). Observations are provided every second year from 1961 to 2001. The time series are interpolated over the years, and extracted by the last known observation.

BD: Benefit duration

Benefit duration is a measure on how long the benefits last when being unemployed. The ratio is calculated by the time series described under benefit replacement rates and equation (A5).

$$BDj_{it} = \alpha \frac{brrja2_{it}}{brrja1_{it}} + (1-\alpha)\frac{brrja4_{it}}{brrja1_{it}} \tag{A5}$$

where $\alpha = 0.6$, $j = \{67, 100\}$, $i = 1, 2...20$ and $t = 1960, 1961...2007$. $brrja1_{it}$ is the benefit replacement rate in year 1, $brrja2_{it}$ is the benefit replacement rate in year 2 and 3, and finally $brrja4_{it}$ is benefit replacement rate in year 4 and 5. $\alpha = 0.6$ gives more weight to the second and third year as compared to the fourth and fifth year. The index is calculated for both employment situations, i.e. $j = 67$ percent and $j = 100$ percent of average earnings.

The average of $bd67_{it}$ and $bd100_{it}$ is used as an indicator of benefit duration, i.e. BD_{it}.

If benefit duration stops after one year, then brr67a2 = brr67a4 = 0, and BD67 = 0. If benefit provision is constant over the years, then brr67a1 = brr67a2 = brr67a4, and BD67 = 1. However, some countries increase payments over time and the value of benefit duration is above one.

Dudnet: Union density

Union density rates are constructed using the number of union memberships divided by the number of employed.

The main data source is Visser (2009), where they have mainly calculated the trade union density index based on surveys. When data were unavailable, they have used administrative data adjusted for non-active and self-employed members. The database Nickell (2006) contains additional information for Sweden before 1975 and Ireland in 1960. The time series for Sweden in the latter source is interpolated, are this growth rate is then used to prolong the original time series from Visser (2009).

The interaction terms between union density and coordination are prolonged by the last known observation for these countries.

CO: Coordination of wage setting

The index for coordination of wage setting describes the coordination level in the wage setting. The index ranges from 1 to 5, and the most coordinated countries have index equal to 5:

The main source is OECD (2004), see table 3.5. The frequency for observations is five-year intervals over the period 1970-2000. The years are interpolated between means, i.e.

between 1972-1977, and with 1970 and 2000 equal to the first and last five-year intervals. In the period 1960 to 1970 the observations are prolonged backwards by the last known observation for all countries. The same procedure is used to extend the time series until 2007.

EPL: Employment protection

The time series for employment protection measures the strictness of the employment protection for the employer. The overall measure for employment protection is measured on a scale from 0 to 5. Strictness is increasing in scale. Some other measures only measure the employment protection for regular- or temporary employment.

The time series for employment protection is provided by OECD (2004) for the period 1985 to 2004. The time series are based on the point observations in Annex 2.A2. 6. Note, however, that OECD (2004) claims that judgement is made when constructing the time series. This implies that time series for employment protection not are only a linear interpolation between the point observations. The measure of employment protection refers to the protection of overall employment (*epl*).

Before 1985, the time series are prolonged backwards using data from Belot and van Ours (2004). The source contains data in five-year intervals, but the data are here used annually by interpolation between the means of the observation points. The percentage change is used to prolong the time series in OECD (2004), by equation (A6):

$$Y_t = Y_{t+1} * \frac{X_t}{X_{t+1}} where Y = \{epl, eplr, eplt\} and X = \{EPLBO\_bo, EP\_bo\} \tag{A6}$$

Portugal and Spain are prolonged backwards by the last known observation in the period 1960 to 1984. The United States is prolonged backwards by the last known observation in the period 1960 to 1982.

4.B Extra

Table B1 presents the complete results of the models in table B1, including the estimates for the labour market variables. Thus, model 1 is reestimation of the unrestricted unemployment equation of Nymoen and Sparrman (2010), including dummies for large outliers. Model 2 includes government purchases and the indicator for the export market, and in model 3, the dummies for large outliers are omitted.

For comparison, table B1 also includes two models more, 2i and 3i. Comparing models 2 and 2i shows the effect of omitting the large outliers, which is a fairly large increase in the coefficient values of government purchases and the export market. Model 3i extends model 3 by including year dummies. We observe that the coefficients for the export market become much smaller, reflecting considerable co-movement of the export markets for all countries. In contrast, the coefficient for the change in government purchases is not affected, presumably because any comovement in government purchases across countries is not linked to comovement in unemployment.

Table B1: Estimation of equation (1) with the change government purchases and export markets.

	Model 1			Model 2			Model 2 i			Model 3			Model 3 i [a]		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
Unemployment previous period	1.38	0.03	0.00	1.29	0.03	0.00	1.28	0.04	0.00	1.28	0.04	0.00	1.26	0.04	0.00
Unemployment two years ago	-0.52	0.05	0.00	-0.41	0.05	0.00	-0.40	0.05	0.00	-0.39	0.05	0.00	-0.40	0.05	0.00
Unemployment three years ago	0.06	0.03	0.04	0.01	0.03	0.70	-0.01	0.03	0.86	-0.01	0.03	0.82	-0.00	0.03	0.96
Institutions:															
Employment protection (EPL), 1st diff. previous period	0.12	0.24	0.61	0.09	0.22	0.69	0.04	0.27	0.87	0.04	0.27	0.87	-0.25	0.25	0.32
EPL, two years ago	0.14	0.07	0.05	0.14	0.07	0.03	0.16	0.08	0.05	0.15	0.08	0.05	0.04	0.08	0.60
Benefit replacement ratio (BRR), 1st diff. previous period	-0.89	0.81	0.27	-0.19	0.78	0.81	1.16	0.94	0.22	0.98	0.93	0.29	0.29	0.88	0.74
BRR, two periods ago	0.63	0.24	0.01	0.61	0.25	0.01	0.97	0.29	0.00	0.92	0.29	0.00	0.70	0.27	0.01
Benefit duration (BD), 1st diff. previous period	-0.51	0.54	0.35	-0.53	0.53	0.31	-0.77	0.63	0.22	-0.88	0.62	0.16	-0.58	0.59	0.33
BD, two periods ago	0.00	0.17	1.00	-0.00	0.17	0.98	-0.13	0.20	0.51	-0.14	0.20	0.47	-0.16	0.19	0.41
Interaction - BRR and BD 1st diff. previous period	-2.59	2.03	0.20	0.02	1.99	0.99	4.10	2.37	0.08	3.74	2.31	0.11	3.05	2.16	0.16
Interaction - BRR and BD two periods ago	1.17	0.63	0.06	1.12	0.67	0.09	1.96	0.80	0.01	1.99	0.80	0.01	2.04	0.76	0.01
Interaction - CO and UDNET 1st diff. previous period	-4.16	2.15	0.05	-4.61	2.03	0.02	-2.88	2.44	0.24	-2.91	2.42	0.23	-2.03	2.28	0.37
Interaction - CO and UDNET two periods ago	-0.77	0.46	0.10	-0.67	0.44	0.13	-1.07	0.53	0.04	-1.09	0.52	0.04	-0.90	0.48	0.06
Interaction - CO and TAX 1st diff. previous period	-8.02	2.48	0.00	-7.73	2.35	0.00	-11.99	2.81	0.00	-11.71	2.80	0.00	-8.43	2.67	0.00
Interaction - CO and TAX two periods ago	-0.18	0.81	0.82	-0.84	0.78	0.29	-1.03	0.94	0.27	-0.82	0.93	0.38	-0.26	0.87	0.77
Union density (UDNET), 1st diff. previous period	0.43	2.02	0.83	-0.91	1.95	0.64	-1.21	2.35	0.61	-0.67	2.30	0.77	1.36	2.23	0.54
UDNET, two periods ago	0.26	0.28	0.35	0.02	0.29	0.96	0.42	0.34	0.22	0.53	0.34	0.11	0.34	0.37	0.36
Coordination (CO), 1st diff. previous period	0.12	0.17	0.46	0.19	0.16	0.23	-0.34	0.19	0.07	-0.36	0.19	0.05	-0.32	0.18	0.08
CO, two periods ago	-0.01	0.04	0.79	-0.01	0.04	0.84	-0.07	0.05	0.15	-0.08	0.05	0.10	-0.06	0.04	0.15
Tax level (TAX), 1st diff. previous period	-0.29	1.52	0.85	1.32	1.46	0.36	2.04	1.75	0.24	1.95	1.74	0.26	2.96	1.67	0.08
TAX, two periods ago	0.48	0.53	0.37	0.57	0.53	0.28	1.41	0.64	0.03	1.48	0.63	0.02	1.19	0.68	0.08
Large outlier[b]	0.94	0.05	0.00	0.82	0.04	0.00									
Export market, 1st diff. (ΔXM_t)				-0.39	0.05	0.00	-0.53	0.06	0.00	-0.53	0.06	0.00	-0.21	0.10	0.04
Export market, prev. period (XM_{t-1})				0.12	0.06	0.03	0.17	0.07	0.01	0.17	0.07	0.01	0.13	0.10	0.19
Export market, 1st diff. prev. period (ΔXM_{t-1})				-0.23	0.06	0.00	-0.31	0.07	0.00	-0.30	0.07	0.00	-0.31	0.10	0.00
Demand components:															
Change govt. purchases, 1st diff. (ΔdG_t)				-0.14	0.03	0.00	-0.20	0.04	0.00						
Change govt. purchases, prev. period (dG_{t-1})				-0.21	0.05	0.00	-0.22	0.06	0.00						
Change govt. purchases, 1st diff. prev. period (ΔdG_{t-1})				0.03	0.03	0.31	0.02	0.04	0.60						
Change govt. purchases, (dG_t)										-0.19	0.04	0.00	-0.20	0.04	0.00
Obs = Country*Average groups	837	20	41.9	794	20	39.7	794	20	39.7	801	20	40.0	801	20	40.0
Standard deviation of residuals	0.58			0.54			0.65			0.65			0.60		
χ^2 of all the exogenous variables.[c]	531.93	(0.00)		766.19	(0.00)		286.23	(0.00)		287.57	(0.00)		103.97	(0.00)	
χ^2 of dummy, fiscal policy and exports.[c]				635.76	(0.00)										
χ^2 of policy and exports.[c]				150.38	(0.00)		195.92	(0.00)		195.35	(0.00)		52.92	(0.00)	
1st order autocorrelation[c]	0.37	(0.71)		0.29	(0.77)		0.83	(0.41)		0.65	(0.52)		1.20	(0.23)	
2nd order autocorrelation[c]	-1.71	(0.09)		-2.00	(0.04)		0.28	(0.78)		0.15	(0.88)		0.79	(0.43)	

Estimation method: Fixed effect coeffcients estimate, standard errors from GLS (xtgls without options) is used in all the regressions.

a) With time dummies.

b) Break by large outlier approach.

c) Numbers in parenthesis are p-values for the relevant null.

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